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The Pass-Through of Exchange Rate Changes to Prices in the Euro Area: An Empirical Investigation

Nidhaleddine Ben Cheikh

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sous le sceau de l'Université Européenne de Bretagne

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Ecole doctorale Sciences de l'Homme et de la Société

présentée par

Nidhaleddine Ben Cheikh

préparée à l'unité de recherche CREM-UMR 6211
Centre de Recherche en Economie et Management

**The Pass-Through
of Exchange Rate
Changes to Prices
in the Euro Area:
An Empirical
Investigation**

Thèse soutenue à Rennes

Le 14 Octobre 2013

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L'Université de Rennes 1 n'entend donner aucune approbation ni improbation aux opinions émises dans cette thèse. Ces opinions doivent être considérées comme propres à leur auteur.

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INTRODUCTION AND OVERVIEW

“The secret of empirical work is to define your hypothesis so that failure to find significant results can be interpreted as support.”

— Jeffrey A. Frankel

1. Background and Rationale of the Research

What is Exchange Rate Pass-Through?

The textbook definition of the Exchange Rate Pass-Through (henceforth ERPT), as reported in GOLDBERG and KNETTER (1997, p. 1248), is “the percentage change in local currency import prices resulting from a one percent change in the exchange rate between the exporting and importing countries”. Although this concept is traditionally related to the responsiveness of the *prices of imported goods* to movements in the nominal exchange rate, the definition has evolved over time to include other types of prices, notably producer prices and consumer prices. Thus, the ERPT can be seen more broadly as the change in domestic prices (import prices, producer prices and consumer prices) that can be attributed to the change in the nominal exchange rate. Also, we can add that, as is common in the literature, we call the “first-stage pass-through” as the transmission of the exchange rate changes to import prices, and the “second-stage pass-through” as the responsiveness of consumer prices to changes in import prices. As regards the magnitude of the pass-through, if the effect of the exchange rate changes is fully transmitted to domestic prices, then ERPT is said to be full, or complete. If only a

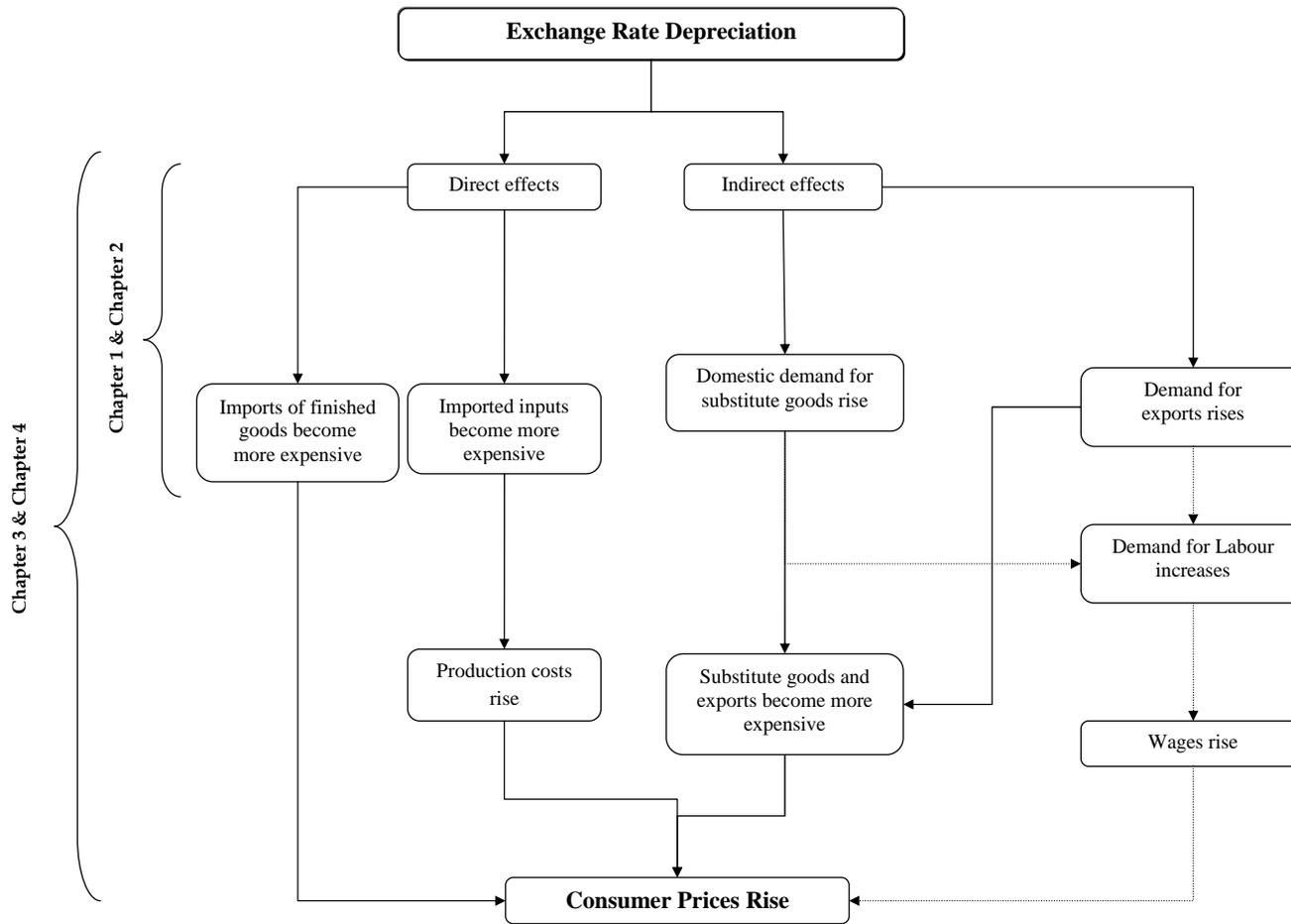
portion of the exchange rate variation is reflected in prices, then ERPT is considered as partial, or incomplete.

In general, the impact of exchange rate fluctuations on domestic prices can be transmitted through *direct* and/or *indirect* channels as can be distinguished in the literature (see Figure 0.1).

The *direct* channel refers to the direct effect that an increase (or decrease) in the external value of a currency has on the price of imported finished goods and imported inputs. On one hand, when the exchange rate depreciates, imported finished goods become more expensive for domestic consumers, and thus domestic consumer prices will in line with the share of imports in the consumption basket. On the other, currency depreciation would entail higher costs of imported inputs leading to higher prices of domestically produced goods, if domestic producers or wholesalers raise their prices in line with the increase in import prices, which can be reflected in consumer prices.

The *indirect* effects of exchange rate changes refers to the competitiveness of goods on international markets through its effect on the aggregate domestic demand and wages. A depreciation of the exchange rate will change the composition of demand, raising both domestic and foreign demand for domestic goods as they become cheaper relative to foreign goods. If the economy is already working at high levels of capacity utilization, the increase in the exports and aggregate demand puts up inflationary pressures on the economy. Also, the increase in the demand for domestic products leads to a higher demand for labour and, potentially, to rising wages, which will in turn be reflected in higher prices. Another important second-round effect which deserves to be mentioned is related to nominal wage rigidity in the short run. When domestic prices are rising, real wages will decrease and output will increase. To the extent that real wages will be regain their original level over time, production costs increase, the overall price level increases and output falls. Thus, in the end the exchange rate depreciation leaves a permanent increase in the price level with only a temporary increase in output.

Figure 0.1: Pass-through from an exchange rate depreciation to consumer prices



Source: LAFLÈCHE (1996).

Chapter 3 & Chapter 4

Chapter 1 & Chapter 2

Why do we care about Exchange Rate Pass-Through?

A number of studies have been motivated to examine more closely the underlying relationship between the exchange rate and prices. Thorough knowledge of the degree of and underlying behavior behind pass-through is of particular importance for several policy issues, including for the design of monetary policy, adjustment in trade balances, the international transmission of shocks and the optimal choice of exchange rate regime.

As import prices are a principal channel through which movements in the exchange rate affect domestic prices and hence also the variability of inflation and output, these considerations would ultimately have important implications for the appropriate stance of monetary policy. If the inflationary effects of exchange rate changes are large, the central bankers will have to implement monetary policies that could offset the inflationary consequences of exchange rate changes. Policymakers must be able to gauge how large these effects are likely to be, in order to determine the size and persistence of underlying inflation pressures and any monetary policy responses that might be required to deal with them. Also, as is well-known the successful implementation of monetary policy presupposes that central bankers have not only a good understanding of inflation dynamics, but that they are also relatively successful at predicting the future path of inflation. Thus, the monetary authorities' forecasts of the future path of inflation must factor in the changing behaviour in the ERPT. A frequently cited example is the decline of the sensitivity of domestic prices to exchange rate movements in the last two decades. If inflation forecasts are based on estimates of ERPT that do not take into account such a decline, these forecasts could be overestimating the effects of changes in the exchange rate on inflation. Besides, it is important to note that adoption of inflation targets by central banks in many countries would their concern about the size and speed of the ERPT into domestic prices.

In addition, the pass-through of exchange rate it is a key input for determining the path of external adjustment. The extent of ERPT will influence domestic demand for real imports and thus contribute to the adjustment (or non-adjustment) of real domestic trade balance. When the degree of pass-through to tradable prices is found to be high, the exchange rate changes will affect the relative prices of tradables and non-tradables, so that the adjustment in the current account balance will be relatively prompt. For instance,

imported goods become expensive, if pass-through is high, so that expenditure switching from imports to domestic goods will occur and external balances will be corrected in several months. On the other hand, If prices respond sluggishly to changes in exchange rates and if trade flows respond slowly to relative price changes, this does not help external adjustment of the economy. In other words, a low degree of pass-through would make it possible for trade flows to remain relatively insensitive to changes in exchange rates. For example, currency depreciation would not reduce imports or promote exports to correct the external imbalances.

Furthermore, the degree to which exchange rate fluctuations are reflected in prices also matters for the international transmission of monetary shocks. For example, a depreciation of the domestic currency which is a result of a positive monetary shock generates an expenditure-switching effect, shifting world demand away from foreign goods towards domestic goods. Consequently, output rises in the country where the depreciation has occurred and falls abroad. Thus, we can say that monetary disturbances will tend to generate negative comovements of output across countries, if the extent pass-through is high. However, if the degree of pass-through is sufficiently weak, this ordering is reversed, and the cross-country correlation of output becomes positive. This outcome has an important implication. As monetary policy shocks are important in explaining business cycles, the recent decline in pass-through (as reported in the bulk of the empirical literature) would enhance the comovement of outputs across countries and business cycles will become more synchronized.

Finally, the effects of exchange rates on prices can determine the choice of a fixed or flexible exchange rate regime. It is known that exchange rate flexibility facilitates allows immediate relative price adjustment in response to real shocks. The adjustment of relative prices generates an expenditure-switching effect between home and foreign goods that partly offsets the initial effect of the shock. This argument supposes that domestic currency prices of imported goods respond to movements in nominal exchange rates. If the degree of exchange rate pass-through is low, the expenditure-switching effects will be weak, thus limiting the short-run adjustment role of nominal exchange rates and, hence, a flexible exchange rate will not offer any advantage. However, when pass-through to import prices is complete, i.e. import prices respond one-to-one to exchange rate changes, a flexible exchange rate regime is desirable because it allows

relative price adjustments to occur. However, under null pass-through, i.e. import prices do not respond at all to currency movements, a flexible exchange rate will not offer any advantage. The optimal policy involves fixing the nominal exchange rate because since flexible exchange rates cannot achieve the optimal relative price adjustment.

To conclude about the policy relevance of the pass-through, it is crucial to distinguish between the ERPT to import prices and the ERPT to consumer prices. If all domestic prices respond to the nominal exchange rate depreciation one-to-one, i.e. complete pass-through not only to import prices but to consumer prices, then any export competitiveness gained from nominal depreciation would be cancelled out the inflationary impact of the domestic currency's fall. As a result, the real exchange rate would not change at all since combination of nominal depreciation and high inflation leaves the export competitiveness unchanged. Thereby, from the viewpoint of using the exchange rate changes as an instrument for correcting the external imbalances, a higher pass-through to import prices is desirable, while a greater pass-through to consumers prices, raising all price levels, is harmful. It is important that monetary policy is conducted with knowledge of this distinction. Thus, one of the key questions for policy makers is: *How much of an exchange rate change is "passed through" to import prices and to overall consumer prices?*

Why focusing on the Euro Area case?

The issue of ERPT is of particular interest for a monetary union such as the euro area (EA). For the European countries, forgoing their local currencies to join a monetary union has posed a significant challenge, since a country adopting the euro cedes its monetary policy to the European Central Bank (ECB), and no longer has the option of using monetary policy to respond to local conditions. The impact of the monetary policy decisions on the single currency may induce different effects on expenditure switching and price level movements regarding the extent of the transmission of exchange rate changes to domestic prices. Thereby, a common exchange rate movement, in the absence of a national monetary policy, may have differential impact on different EA member states, leading notably to possible divergence in inflation rates.

Nevertheless, as the main objective of the monetary policy of the ECB is to achieve medium-term price stability for the euro zone aggregate, this may be seen as a sign of increasing credibility of the monetary regime for the countries joined the euro area. This is true especially for countries with historically higher levels of inflation, such as Greece, Italy, Portugal and Spain. As argued in the ERPT literature, a more stable monetary policy conditions with credible and anti-inflationary regime tend to reduce the degree to which the currency changes are transmitted to domestic prices. One of the first to put forward this argument was TAYLOR (2000). The author postulates that the prevailing of environment low-inflation regime, which serves to reduce the perceived persistence of cost shocks, would lower the degree of ERPT. Hence, one can think that the start of Stage III of the European Monetary Union (EMU) in January 1999 would affect the behaviour of the pass-through across the EA countries.

In addition to the change in macroeconomic environment, there are other reasons which can explain why the rate of pass-through might have changed as a result of the introduction of the euro. The proportion of trade exposed to exchange rate movements has diminished after the adoption of the single currency, leading to change in the relative degree of openness in the monetary union members. As explained by DORNBUSCH (1987), pass-through may be higher if the exporters are large in number relative to the presence of local competitors. However, the advent of the euro has reduced the market power of foreign firms relative to their domestic counterparts, this changing in the competitive conditions may lead to a decline in the responsiveness of import prices. Furthermore, the choice of the currency of invoicing for trade flows would be affected following 1999. It is expected that the share of imports being denominated in the euro would increase for the whole EA countries. DEVEREUX, ENGEL, and TILLE (2003) argued that as the euro become a well established currency, it will be a “vehicle currency” which competes with the US dollar, favouring the expansion of the euro as a currency of denomination of imports across member countries. To the extent that the single currency is chosen for transaction invoicing by foreign firms, imports prices in the euro zone become more insulated from exchange rate fluctuations. As a result, the transmission from exchange rates to import prices tend to be lower as imported goods being more local-currency priced. The consequence of increasing share of imports denominated in the euro has been explicitly expressed by the ECB: “...increasing use of the euro as a

*payment/vehicle and pricing/quotation currency could have two effects. First, it could make the euro area Harmonized Index of Consumer Prices (HICP) less sensitive, in the short run, to US dollar exchange rate movements...Second,...the short-term effects of exchange rate changes on the goods and services trade balance should generally be reduced.”*¹

Otherwise, since the start of the monetary union, the EA countries have been subject to substantial fluctuations in the exchange rate of the euro. During the first three years of his existence, the euro experienced a large depreciation of roughly 45 percent against the U.S. dollar and about 25 percent on a trade-weighted basis. This extensive depreciation was followed by roughly the same magnitude of appreciation between 2002 and 2004 (see Figure 0.2). These wide swings have raised concerns that it might lead to higher inflation variability. Especially, the euro weakness may likely raise the cost of imports and producer prices, which can feed into higher consumer prices. The concern about the exchange rate affecting price stability has been clearly expressed by the monetary authority in the EA: *“Developments in the exchange rate of the euro are becoming a cause for concern with regard to future price stability...Given both the magnitude and the duration of this development, import prices can be expected to rise further, thereby increasing the risk that upward pressures on consumer price inflation might materialize in the medium term.”*² Thus, “imported” inflation remains a threat for the monetary union and may impact differently the EA member countries, depending on their relative different exposures to extra-EA trade. For instance, a member country with large imports from a non-EA country will experience different inflationary pressures if the euro depreciates as compared to a member country that conducts all its trade with other member countries.

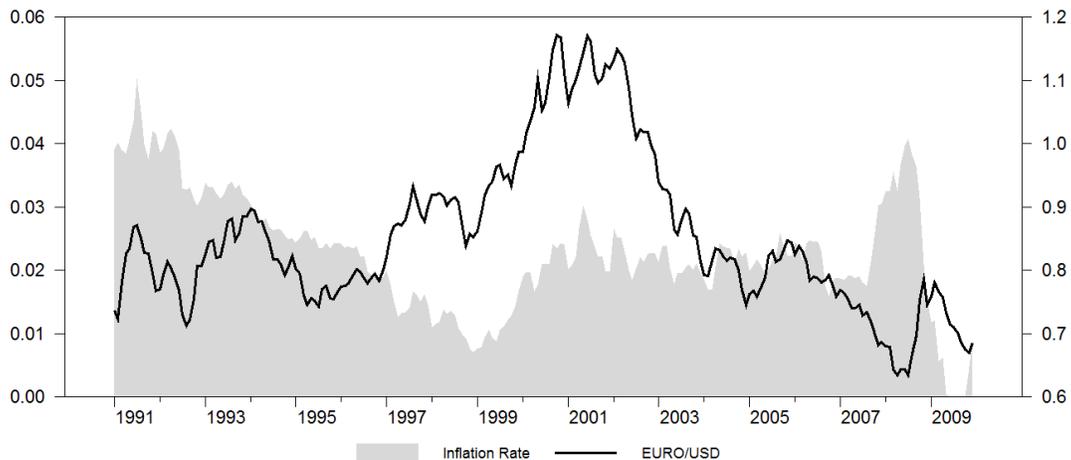
The potential effects of these dramatic exchange rate movements on inflation and trade have taken on renewed interest and significance, and forced the ECB to take them into account when making monetary policy decisions. As a matter of fact, the ECB concerns about the economic impact of past euro appreciation have figured prominently. The European monetary authorities cited the inflationary effects of a lower value of the euro as a factor behind its tightening of monetary policy in 2000 and the disinflationary

¹The Monthly Bulletin of the ECB, August 1999.

²Willem F. Duisenberg, President of the ECB, Press conference, 3 February 2000.

effects of a strong euro as a factor behind the loosening in 2003.³ Overall, the different arguments given above constitute a motivation for us to study the extent of ERPT in the EA countries, and to answer the question of whether the launch of the euro is a watershed event in this respect.

Figure 0.2: Exchange Rate and inflation in the euro area



Source: International Financial Statistics of IMF.

Incomplete and declining ERPT: a “Macro” or “Micro” phenomenon?

There has been a large body of empirical and theoretical literature on the relationship between exchange rate and prices. This is not surprising since the analysis have followed various paths, starting with debates on the validity of the law of one price and purchasing power parity, followed by explaining the role of market power and price discrimination in international markets, and, more recently, the debates has evolved by stressing on the relevance of macroeconomic policies in determining the pass-through. An extensive survey of the literature was provided by GOLDBERG and KNETTER (1997). As reported by the authors, a search in the EconLit database of the words “Law of One Price”, “Purchasing Power Parity”, “Exchange-Rate Pass-Through” and “Pricing-to-Market”

³See the statements given by the ECB in connection with Council monetary policy decisions between February and July 2000.

yielded nearly 700 entries. In our case, when restricting the search to only the key words “exchange rate pass-through”, Econlit database returns a total of more than 800 entries at the end of 2012.⁴

Acknowledging this massive literature, here we only focus on the main findings and features of the ERPT literature. Indeed, the most common findings is that the effects of exchange rate changes on the different domestic prices - import prices, producer prices and consumer prices - are incomplete and has declined markedly in recent years. These regularities appear to be valid especially for the industrialized countries. In some empirical studies, the import-price pass-through, which is expected to be higher than in producer and consumer prices, is found to be incomplete even in the long-run. In spite of the overall consensus, there is a substantial debate about the conditions that lead to low and declining ERPT. In other words, the literature has had a hard time to pin down with certainty the source of the decline and incomplete degree of pass-through. As a matter of fact, this phenomenon has both macro- and microeconomic aspects as discussed in CAMPA and GOLDBERG (2005), but the literature is not conclusive about the most important factors, i.e. macro or micro factors. Consequently, the vast ERPT literature can be divided them into two strands. The first strand of literature follows an approach inspired by the industrial organization literature and focuses on the role of market structure and foreign firms’ pricing behavior to explain the incomplete ERPT. In their seminal papers, DORNBUSCH (1987) and KRUGMAN (1987) justifies incomplete pass-through as arising from foreign firms that operate in a market characterized by imperfect competition and adjust their mark-up to maintain a stable market share in the domestic economy, which can drive less than one-to-one transmission of exchange rate. This exchange rate induced mark-up adjustment is usually referred in the ERPT literature as “pricing-to-market” (PTM) strategy. We note that empirical papers in this strand of literature are industry or product specific studies, i.e. a disaggregated data of different products or industries on the micro level are used. Also, most of these studies focus on the pass-through to export or import prices, neglecting the pass-through to other prices such producer and consumer prices. In the second strand of literature, the incomplete or decline ERPT is rather macroeconomic phenomenon which is related in particular to the inflationary and exchange rate regimes. This category of studies highlights the role of

⁴Econlit and Business Source Complete databases together yield more than 1400 entries for the terms “exchange rate pass-through”.

macroeconomic environment and in particular the monetary policy regime. It stipulates that countries with low relative exchange rate variability or stable monetary policies are more likely to have their currencies chosen for transaction invoicing, and hence more likely to have low pass-through to domestic prices. We note that the second strand studies the effects of exchange rate pass-through on the macro level using aggregate price measures. As they aim at providing evidence that is more relevant for macroeconomic policy, pass-through of exchange rate changes to import, producer and consumer price are all of interest. So they follow the broad definition of pass-through and measure the pass-through rates of exchange rate changes to not only import price, but also producer and consumer prices. It is worth noting that the majority of the ERPT papers are micro-level studies, there has been a revival of interest for macroeconomic factors in the recent years, since studying of a particular product or industry limit the ability for international comparisons due to lack of data availability.

Moreover, the distinction between macro- and microeconomic factors is very important since they point to substantially different implications in policy terms. If pass-through is a “macro” phenomenon which is directly associated with monetary policy, such as inflation or exchange rate volatility, this implies that a given decline in pass-through, may not necessarily be a permanent phenomenon because it may dissipate if monetary policy becomes more accommodative. In contrast, if ERPT is related to more structural factors, such as the industry composition of trade, economic policy is less capable to deal with. Also, some micro factors may lead to different conclusions. For instance, the role of product differentiation is actually ambiguous as two different effects may cancel out: on one hand, more differentiated goods may be characterized by higher market power and therefore higher pass-through (see e.g YANG, 1997); on the other hand, more differentiated products may be characterized by higher markups, hence higher scope for PTM and therefore lower ERPT (see e.g CAMPA and GOLDBERG, 2005). Thus, the source of incomplete or declining ERPT is a quite relevant issue, as they yield very different policy implications.

Besides, we can add that empirical literature has often reported notable cross-country differences in the rates of pass-through. To give an idea on the variability of ERPT across countries, results of CAMPA and GOLDBERG (2005) for import-price pass-through and GAGNON and IHRIG (2004) for consumer-price pass-through are displayed

in Table 0.1 and Table 0.2 respectively. In these studies, the estimated ERPT elasticities are representative of those found in the broader literature. In fact, the macro- and micro factors listed above would explain why the exchange rate fluctuations are differently transmitted to domestic prices. For 23 OECD countries, CAMPA and GOLDBERG (2005) found that micro determinants, notably the trade composition of imports, are by far more important than macro determinants, such as inflation environment, in explaining the decline of ERPT to import prices. In other words, the authors explained that a shift in the composition of the imports towards sectors with lower degrees of pass-through, such as the manufacturing sector where more differentiated goods are produced and hence PTM strategy is more frequent, explains most of the decline in pass-through to import prices.

Table 0.1: ERPT to import prices in CAMPA and GOLDBERG (2005)

Country	ERPT Elasticities	
	Short-run	Long-run
Australia	0.56* [#]	0.67* [#]
Austria	0.21 [#]	0.10
Belgium	0.21 [#]	0.68
Canada	0.75* [#]	0.65* [#]
Czech Republic	0.39* [#]	0.60*
Denmark	0.43* [#]	0.82*
Finland	0.55*	0.77*
France	0.53* [#]	0.98*
Germany	0.55* [#]	0.80*
Hungary	0.51* [#]	0.77*
Ireland	0.16 [#]	0.06
Italy	0.35* [#]	0.35 [#]
Japan	0.43* [#]	1.13*
Netherlands	0.79* [#]	0.84*
New Zealand	0.22* [#]	0.22 [#]
Norway	0.40* [#]	0.63*
Poland	0.56* [#]	0.78*
Portugal	0.63* [#]	1.08*
Spain	0.68* [#]	0.70*
Sweden	0.48* [#]	0.38* [#]
Switzerland	0.68* [#]	0.93*
United Kingdom	0.36* [#]	0.36* [#]
United States	0.23* [#]	0.32* [#]
Average	0.46	0.64

Note: *^(#) implies that ERPT coefficient is significantly different from 0 (1) at the 5% level. ERPT elasticities are estimated using quarterly data over 1975-2003. Short-run ERPT is defined as the impact of exchange rate within one quarter, while long-run ERPT is the effect of exchange rate after one year.

Table 0.2: Long-run ERPT to consumer prices in GAGNON and IHRIG (2004)

Country	Long-run ERPT Elasticities		
	Entire sample	First sample	Second sample
Australia	0.14 (0.07)	0.09 (0.08)	0.01 (0.04)
Austria	0.11 (0.07)	0.06 (0.10)	0.04 (0.02)
Belgium	0.20 (0.08)	0.21 (0.09)	0.02 (0.02)
Canada	0.37 (0.11)	0.30 (0.14)	0.04 (0.06)
Finland	0.01 (0.14)	-0.11 (0.21)	0.00 (0.03)
France	0.23 (0.12)	0.17 (0.07)	0.01 (0.03)
Germany	0.11 (0.04)	-0.13 (0.11)	0.12 (0.03)
Greece	0.52 (0.11)	0.28 (0.12)	0.27 (0.21)
Ireland	0.29 (0.09)	0.18 (0.11)	0.06 (0.04)
Italy	0.37 (0.12)	0.33 (0.09)	0.08 (0.06)
Japan	0.21 (0.09)	0.26 (0.12)	0.02 (0.02)
Netherlands	0.16 (0.07)	0.08 (0.11)	0.06 (0.03)
New Zealand	0.42 (0.10)	0.29 (0.09)	0.01 (0.05)
Norway	0.28 (0.15)	0.11 (0.17)	-0.05 (0.06)
Portugal	0.43 (0.08)	0.37 (0.08)	0.17 (0.16)
Spain	0.18 (0.09)	0.14 (0.07)	0.03 (0.03)
Sweden	0.02 (0.07)	0.05 (0.05)	0.02 (0.02)
Switzerland	0.15 (0.09)	0.18 (0.14)	0.07 (0.08)
United Kingdom	0.15 (0.05)	0.18 (0.08)	0.08 (0.05)
United States	0.27 (0.12)	0.19 (0.36)	0.03 (0.06)
Average	0.23	0.16	0.05
Inflation targeters	0.22	0.18	0.03
Non-targeters	0.23	0.15	0.06

Note: Numbers in parentheses are standard errors. ERPT coefficients for the entire time sample are estimated using quarterly data over 1971-2003. Sub-periods estimations are different for each country and are based on the level of inflation.

As regards the consumer-price pass-through, GAGNON and IHRIG (2004) found a substantial role of inflation regime in explaining the lowering ERPT for 20 industrial countries in the recent years. The authors create two subsamples, with sample break dates chosen independently for each country, based on the observed behaviour of inflation. The first subsample period is a period of relatively high and variable inflation, whereas the second subsample has lower and more stable inflation. As showed in Table 0.2, the extent of pass-through differs strongly between the two subsamples. Especially for countries that have adopted inflation targeting the reduction is more pronounced. Since there is no consensus, the source of the observed decline in ERPT is still an open issue which needs a more thorough analysis.

Could methodology account for uncertainty in estimates?

In a comprehensive survey of the ERPT literature, MENON (1995) summarized the results of 43 papers and revealed some shortcomings of previous empirical pass-through studies. More specifically, the author suggested that for a given country or industry the estimated ERPT coefficients are found to be different across different studies. The author attributes these divergences to heterogeneity of methodologies, model specification and variable construction rather than from different time periods studied. This leaves no clear picture of the importance of exchange rate changes for domestic economic conditions. In the recent years, there has been some empirical work on pass-through that tried to improve the deficiencies of earlier studies that were identified by MENON (1995). Thus, here we report the main criticism of the earlier empirical studies on the degree of pass-through and discuss the alternative approaches that were suggested recently to obtain reliable estimates.

In fact, the earlier empirical literature on ERPT resorts to typical single-equation approach by employing ordinary least squares (OLS) regressions. Justified by an underlying partial equilibrium model, the empirical specifications in this literature assume the domestic prices to respond to an exogenous movement in the nominal exchange rate. Taking the process of the exchange rate as exogenous in the economy and ignoring its potential endogeneity to other variables is in some extent unrealistic. For instance, in the floating exchange rate regime, the exchange rate is one of the endogenous variables that may fluctuate in response to economic policies. Also, according to PPP theory, the relative price levels may drive the exchange rate, then there could be a two-way causality between these variables, and it is more appropriate to adopt an approach that would treat both of them endogenous. As a remedy, the VAR models are proposed to solve endogeneity problem inherent in the single-equation-based methods. This approach allow for system estimation where the endogenous variables are simultaneously determined. Also, the VAR system provide estimated impulse response functions which trace the effects of a shock to one endogenous variable on other variables, allowing us the assessment of ERPT not only within a specific time period, but also its dynamics through time. Furthermore, for nine developed countries, MCCARTHY (2007) has estimated a first-difference VAR model that incorporated prices along the distribution pricing chain, i.e. import prices, producer prices and consumer

prices, in a unifying model, while the previous studies has done it in separate models. Thus, MCCARTHY (2007) framework has the advantage to allow for underlying dynamic interrelations among prices at different stages of distribution and other variables which cannot be done within single-equation method.

Also, another important shortcoming mentioned by MENON (1995) is that the time-series properties of the data, particularly the non-stationarity and cointegration in the data, are not properly taken into account. Failed to find evidence in the data for cointegration, several studies has estimated ERPT models in first differences where the information contained in “levels” variables is lost. Nevertheless, as predicted by the theoretical underpinning of the ERPT mechanism, a long-run or steady-state relationship between the levels of the key variables, i.e. exchange rate and price series, should exist. Thus, using appropriate estimation techniques would help restore a cointegrated equilibrium relationship between the variables in levels. For instance, HÜFNER and SCHRÖDER (2002) suggested to reestimate the pricing chain VAR model of MCCARTHY (2007) a Vector Error Correction Model (VECM) that incorporate the long run relationships among the variables. Due to recent developments in time-series and panel data econometrics, several of the latest studies of exchange rate pass-through explicitly recognize the fact that exchange rate and price series are often non-stationary and may be cointegrated. Within panel data cointegration techniques, DE BANDT, BANERJEE, and KOZLUK (2008) were able to find a strong evidence of cointegrating relationship consistent with the theoretical prediction of a steady state in the ERPT mechanism.

Otherwise, it is worth highlighting that the issue of nonlinearities and asymmetries in the ERPT mechanisms has received little attention even in the recent empirical literature. For example, the question of possible asymmetry of pass-through in appreciation and depreciation periods has been seldom treated in the literature. The number of studies which have investigating for nonlinearities in the context of pass-through is to date relatively scarce, and most of papers assume linearity rather than testing it. The sparse empirical evidence on this area of research has put forth the role of exchange rate movements in generating nonlinearities. According to this literature, mainly, there are two potential sources of pass-through asymmetry. On one hand, asymmetry can arise from the *direction* of exchange rate changes i.e., in response

to currency depreciations and appreciations. On the other hand, the extent of pass-through may also respond asymmetrically to the *magnitude* of exchange rate movements, i.e. depending on whether exchange rate changes are large or small. However, as pointed by MARAZZI et al. (2005), previous studies provide mixed results with no clear support for the existence of important nonlinearities. Although the different factors that may lead to nonlinear mechanism in the pass-through, this issue is routinely disregarded in most of the empirical implementations. As a matter of fact, the issue of nonlinearities/asymmetries requires a careful and relevant econometric analysis. For instance, within nonlinear smooth transition models, NOGUEIRA JR. and LEON-LEDESMA (2008) were able to capture nonlinearities in ERPT with respect to number of macroeconomic variables, namely the inflation rate, the size of exchange rate changes, the output growth and two measures of macroeconomic instability.

Finally, we note that besides econometric studies there is a burgeoning literature dealing with the ERPT in the context of new open-economy macroeconomics (NOEM) models. These latter are based mainly on work by OBSTFELD and ROGOFF (1995) which incorporates nominal rigidities and market imperfections into a dynamic stochastic general equilibrium (DSGE) models for open economies. The main advantage of these models is that, as in VAR models, avoid bias problem due to the endogenous variable, and thus take into account the fact that prices and exchange rate are determined simultaneously. Moreover, the DSGE framework allows analysis of the extent of pass-through conditional on specific shocks (e.g. monetary or productivity shocks.). Rather than assuming exchange rate changes are exogenous shocks that affect prices, the comovement between prices and the exchange rate may depends on the source of the shock. Consequently, the rate of ERPT may be different depending on the nature of shocks hitting the economy. As explained in the NOEM literature, the responsiveness of import prices to exchange rate movements are underestimated in the single-equation regressions, compared to the DSGE models, owing to an econometric bias related to the endogeneity of the exchange rate.

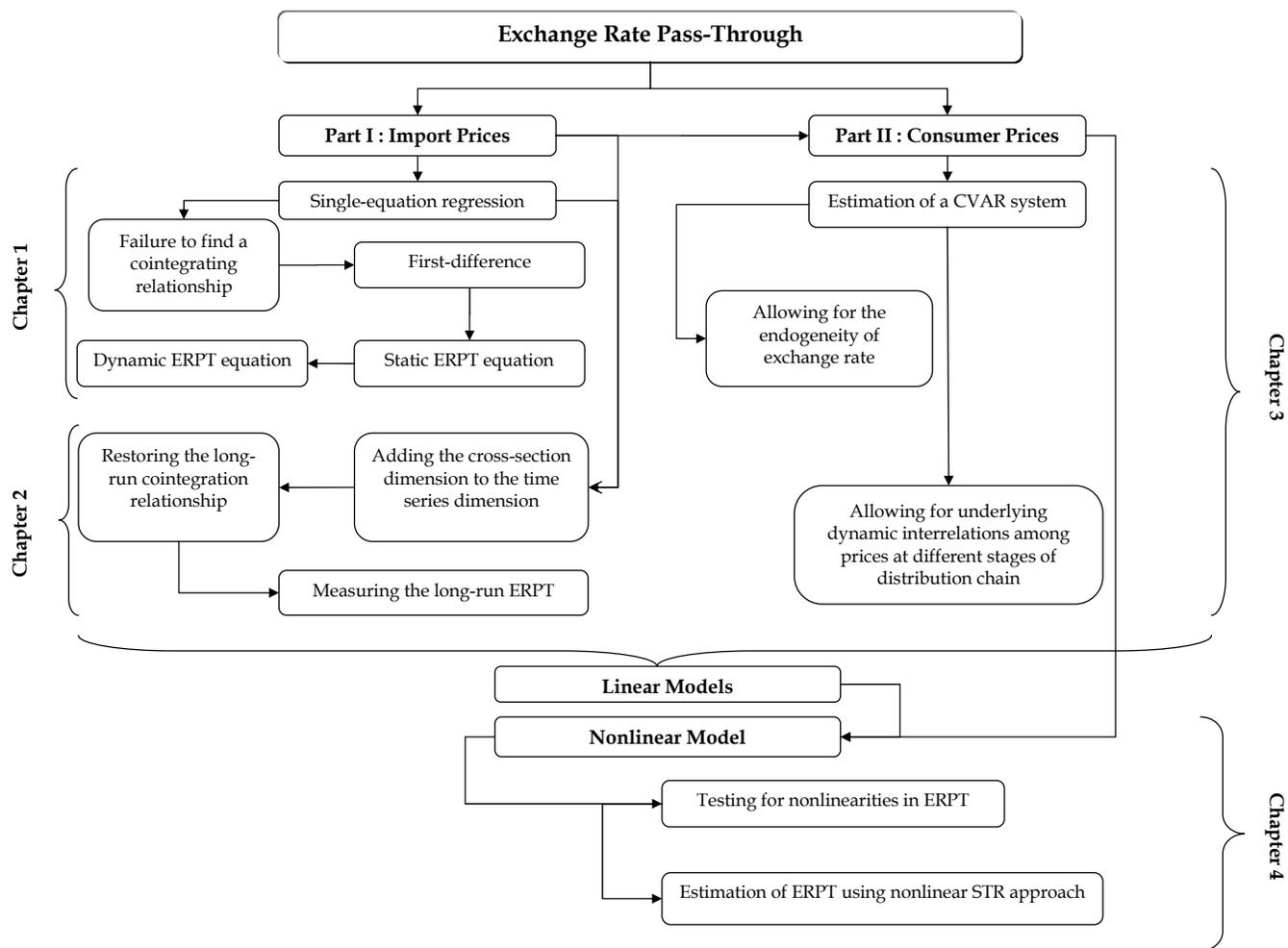
In our exploration of the ERPT issue, however, we use exclusively econometric methods to measure the sensitivity of domestic prices to exchange rate changes, leaving the DSGE models to future research. Thus, an important task for empirical research is

therefore to discriminate between the alternative models and econometric procedures in order to provide valid and reliable ERPT coefficients.

THE OBJECTIVE OF OUR STUDY

Given the wealth of the empirical studies on the degree of ERPT, our research will pursue the following path. First, as many of early studies has dealt with the pass-through from a microeconomic standpoint, our study aims to counter this imbalance by providing a **macroeconomic analysis** of the overall effect of exchange rate changes on domestic prices, an issue which is most relevant for monetary policy. Second, we have explained that the distinction between the import-price pass-through and consumer-price pass-through is crucial since they have different policy implications. It is desirable to have higher ERPT to import prices in order to use the exchange rate changes as an instrument for correcting the trade imbalances, while a lower pass-through to consumer prices is preferred as it avoids inflationary pressure to the domestic economy. Thereby, our thesis will be organized according to this distinction: **the first part** of our analysis will focus on the ‘first-stage pass-through’, i.e. the responsiveness of import prices to exchange rate changes, and in **the second part**, we will examine the transmission of the exchange rate movements via import prices to consumer prices. We can say that our study is concerned primarily with the direct effects of the currency fluctuations (the solid lines in Figure 0.1) and does not explicitly consider the indirect effects via on the aggregate domestic demand and wages (the dotted lines in Figure 0.1). Third, we focus on the countries belonging to the *Eurozone*. The different exchange rate arrangements, in addition to the changing in macroeconomic conditions over time across the EA members states, constitutes a motivation for us to study the extent of ERPT this group of countries, and more specifically to answer the question of whether the launch of the euro is a watershed event that may alter the mechanism of pass-through. Forth, our research analyze the impact of exchange rate movements from an *empirical point of view*, by estimating a wide range of up-to-date econometric methods (dynamic panel data, panel cointegration, CVAR analysis, nonlinear STR models) in order to provide robust measures of the ERPT elasticities as well as to shed further light on the macro determinants of these pass-through coefficients (see Figure 0.3).

Figure 0.3: Empirical methodology for estimating the ERPT



Note: Structure of the thesis from an econometric point of view.

Therefore, our analysis endeavours to illuminate some of the empirical regularities of the ERPT and to contribute to the macroeconomic debate in this regard. The novelties of the thesis are three-fold:

- First, we provide new up-to-date estimates of extent of pass-through for EA countries. There has been a growing interest in the European ERPT in recent years, however, a major drawback was the short time span available since the formation the euro in 1999. Thereby, in our study, we propose an update to ERPT elasticities using longer time period and more observations for the post-EA era.
- Second, contrary to the earlier empirical literature which asserted the prevalence of micro factors, we try to ascertain the role of macroeconomic conditions in influencing the pass-through. As explained above, the source of the decline or incomplete have substantially different implications in policy terms. If pass-through is rather a macro phenomenon which is directly associated with monetary policy, such as inflation environment, this implies that a changing behavior in ERPT is not necessarily a structural phenomenon since it may be solved via macroeconomic policies.
- Third, to give accurate answers to these different questions, up-to-date time series and panel data techniques are provided in this study. Some shortcomings of the empirical literature, like the failure to find a cointegration relationship, assuming exogenous exchange rate or linear ERPT mechanism rather than testing nonlinearity, are resolved by means of wide range of econometric procedures.

2. Overview and main findings of the thesis

Our study provides a detailed examination of ERPT to prices over the last three decades across two categories of prices, namely import prices and consumer prices, thus the thesis will cover two broad themes: the pass-through of exchange rate changes to import prices in Part I (chapters 1 and 2) and the pass-through of exchange rate changes to consumer prices in Part II (chapters 3 and 4).

Part I: Pass-Through to Import Prices

Chapter 1: Measuring Exchange Rate Pass-Through in Euro Area: An Update

There has been a growing interest in the European ERPT in recent years. A common drawback with these studies is the short time span available since the formation the euro in 1999. Thereby, in our study, we propose an update to ERPT elasticities using longer time period and more observations for the post-EA era.

Thereby, the goal of this chapter is to provide new up-to-date estimates of ERPT for 12 EA countries. First, we begin by estimating a static ERPT equation and analyze the main properties of the pass-through elasticities in our sample. This enables us to compare our results with those in the existing empirical literature on the EA, such as CAMPA, GOLDBERG, and GONZÁLEZ-MÍNGUEZ (2005), CAMPA and GONZÁLEZ (2006). These studies used few observations for the monetary union period (post-EA era), thus, their results are updated here. Following this individual estimates exercise, we assess the cross-country differences in our EA sample by investigating whether inflation level and degree of openness of an economy, as potential macro determinants, determine the magnitude of the pass-through. Next, as is typical in the empirical literature, we check the stability of the sensitivity of import prices to exchange rate movements over time. Since the inception of the euro in 1999 constitutes a reduction of shares of imports exposed to exchange rate fluctuations, we test whether the launch of the monetary union in 1999 constitute a break date in the pass-through mechanism. Thereafter, we estimate a dynamic panel data model to provide an aggregate ERPT for the whole EA. The advantage of this framework is to allow us testing the influence of common events (experienced by the EA countries), such as the Exchange Rate Mechanism (ERM) crisis or the formation of the euro, on the responsiveness of import prices to currency changes. Finally, we conduct a sectoral analysis in order to check for the importance of the composition of imports in determining the aggregate pass-through for a country.

Using quarterly data over the period of 1990-2010, we don't find a wide heterogeneity in the degree of pass-through across the 12 EA countries, in contrast to previous empirical works. This is not surprising since previous studies used too few

observations for the EA era, while in our work, the time span for the analysis of the post-EA era is rather long (until the end of 2010). Since the process of monetary union has entailed some convergence towards more stable macroeconomic conditions, it is expected to find a relative low and less dispersed ERPT across EA Member States. Concerning the macro determinants, we found a positive relationship between ERPT and inflation in line with TAYLOR (2000), while no significant role for the degree of openness, measured as the ratio of imports to GDP. Assessing the stability of pass-through elasticities, we find very weak evidence of decline around 1999. However, our results reveal that the pass-through estimates appears to trend down since the beginning of the 1990s. We notice that the observed decline was synchronous to the shift towards reduced inflation regime in our sample of 12 EA countries. It is interesting to note that when we estimate our pass-through equation over 1979:2-1990:2, we point out more pronounced cross-differences in ERPT than recorded over 1990:3-2010:4. There was divergent macroeconomic conditions across EA countries during the 1980s, especially between *peripheral* and *core* EA economies. Thereafter, within the dynamic panel data framework, we confirm the non-significant decline of the import-price sensitivity to exchange rate since the formation of the euro. However, the important role played by inflation environment was confirmed once again. Moreover, our findings suggest that the weakness of the euro during the first three years of the monetary union has raised significantly the extent of pass-through. We pretend that this outcome would explain why the sensitivity of the import prices did not fall since 1999. Finally, using disaggregated import prices data, it appears that the product composition of imports would determine the aggregate ERPT of an economy, and thus, cross-country differences in pass-through rates may be due to heterogeneous industry composition of trade across countries.

Chapter 2: Long-run Exchange Rate Pass-through into Import Prices

This chapter examines the pass-through of exchange rate into import prices in the long-run using recent panel data techniques. Several empirical studies have failed to find evidence of cointegrating relationship in the data. As discussed in panel cointegration literature (see PEDRONI, 1999, 2001, 2004), conventional nonstationary tests have low power in small sample sizes, so adding the cross-section dimension to the time series dimension would increase the power of these tests. Therefore, we propose to use panel

data cointegrating techniques to restore the long-run equilibrium in ERPT relationship. Furthermore, in this study, we attempt to analyze the role of some macroeconomic variables that may account for the cross-country differences in pass-through. Using panel threshold model introduced by HANSEN (1999), we show that our sample of countries can be classified into different groups according to their macroeconomic regimes. This enables us to test the presence of regime-dependence in ERPT mechanism. To the best of our knowledge, this is the first study that applies panel threshold method in this context.

Then, the purpose of this chapter is three-fold: first, we begin by measuring the long-run ERPT into import prices using PEDRONI (2001) methodology by applying FMOLS and DOLS group mean estimators. Initially, this exercise is conducted for 27 OECD countries which include the EA members in order to have more reliable estimates within the panel data framework. Second, we provide insights into the factors underlying cross-country differences in pass-through elasticities. To this end, we explore three macroeconomic determinants, i.e. inflation rate, degree of openness and exchange rate volatility which are potential sources of heterogeneity in ERPT. Due to the important implications of incomplete pass-through for monetary union, in the final part of our analysis, we focus on the case of the euro area by taking a sub-sample of 12 EA countries. Our goal is to assess the behavior of ERPT since the collapse of Breton-Woods era and try to relate it to the change in the inflation environment.

The main results are the following. We first provide a strong evidence of incomplete ERPT in the long-run for our panel 27 OECD countries. Both FMOLS and DOLS estimators show that pass-through elasticity does not exceed 0.70%. When considering individual estimates for our panel of 27 countries, we can note a cross-country difference in the long run ERPT. Especially, there is an evidence of complete pass-through for 5 out of 27 countries, namely, Czech Republic, Italy, Korea, Luxembourg and Poland. Second, when split our sample in different country regimes, we find that countries with higher inflation regime and more exchange rate volatility would experience a higher degree of pass-through. For the degree of openness, our results provide a weak evidence for a positive link between import share and ERPT. When considering the sub-sample of euro area countries, we find a steady decline in the degree of pass-through throughout the different exchange rate arrangements: ERPT

elasticity was close to unity during the *snake-in-the tunnel* period while it is about 0.50% since the formation of the euro area.

Part II: Pass-Through to Consumer Prices

Chapter 3: Pass-Through of Exchange Rate Shocks to Consumer Prices

After focusing on the “first-stage pass-through”, i.e. the sensitivity of import prices to changes in exchange rate movements, in the first two chapters, it is important to examine the overall effect of exchange rate changes on consumer prices, an issue which is more relevant for monetary policy in the euro area. As is well-known, the ERPT to consumer prices involves both first and second-stage pass-through at once, i.e. the transmission of exchange rate changes to import prices, and in turn, the transmission of import prices changes to consumer prices. Thereby, estimating the ERPT to consumer prices would include the effect of exchange-rate movements on both import prices and on other prices in the consumer basket, such as those of domestically-produced goods, services and other non-tradable prices. In order to provide reliable estimates, we need to build a framework that includes different kinds of price indices as well as the nominal exchange rate, allowing us to measure the extent of pass-through at different levels. To achieve this, MCCARTHY (2007) propose a VAR analysis that include all stages of the distribution chain (import, producer and consumer prices) to analyze how exchange rate fluctuations “pass through” to the production process from the import of products to the consumer level. Contrary to the single-equation method, this framework allows for underlying dynamic interrelations among prices at different stages of distribution and other variables of interest. The advantage of simultaneous equation approach allows for potential and highly likely endogeneity between the variables of interest, ignoring such simultaneity would result in simultaneous equation bias. Also, an important drawback regarding some VAR literature, including MCCARTHY (2007), is that the time-series properties of the data - particularly non-stationarity and cointegration issues - was neglected. Then, when a cointegrating relationship is found between variables in levels, it is more appropriate to estimate a Vector Error Correction Models (VECM) that incorporates both short- and long-run dynamics.

Therefore, the objective of this chapter is twofold: On one hand, we seek to remedy some of the shortcomings of the previous studies, by taking into account the non-stationarity and the endogeneity of the variables within a VECM framework. That way, we can analyze the long-run ERPT relationship contained in the cointegrating space. In this exercise, we use a basic VECM model to focus solely on the ERPT to consumer prices. This provides new up-to-date estimates of pass-through for the economies of the euro zone. On the other hand, in the spirit of MCCARTHY (2007), we propose an extended CVAR model that permits to track pass-through from exchange rate fluctuations to each stage of the distribution chain. Several analytical tools are used to explore the impact of exchange rate shocks, namely, impulse response functions, variance decompositions and historical decompositions.

Using quarterly data ranging from 1980:1 to 2010:4, the Johansen cointegration procedure indicates the existence of one cointegrating vectors at least for each EA country of our sample. When measuring the long-run effect of exchange rate changes on consumer prices, we found a wide dispersion of ERPT rates across countries. The degree of ERPT appears to be most prevalent in Portugal and Greece, while the lowest coefficients of long-run ERPT were found in Germany, Finland and France. The distinction between core and peripheral in terms of ERPT is confirmed here. As a natural progression from the cointegration analysis, we carried out impulse response functions analysis. Our results show a higher pass-through to import prices with a complete pass-through detected in roughly half EA countries after one year. These results are relatively large compared to single-equation literature. The magnitude of the pass-through of exchange rate shocks decline along the distribution chain of pricing with the modest effect is recorded with consumer prices. When assessing possible reasons for cross-country differences in the ERPT, inflation level, inflation volatility and exchange rate persistence are the main macroeconomic factors that influencing the degree of pass-through almost along the distribution pricing chain. Next, we have investigated the contribution of external shocks to domestic prices variation by variance decompositions. Results show that contribution is to some extent high in Portugal and Greece. For the latter countries, this would explain why ERPT to consumer is relatively large compared to the other EA members. Finally, using the historical decompositions, we point out that

external factors had important inflationary impacts on inflation since 1999, compared to the pre-EA period.

Chapter 4: Nonlinear Mechanisms of Exchange Rate Pass-through

The issue of nonlinearities is one of the burgeoning topics in the literature of ERPT. In spite of its policy relevance, studies dealing with the nonlinearities in pass-through mechanisms are still relatively scarce. The number of studies which have investigated for nonlinearities in this context is to date relatively scarce, and most of papers assume linearity rather than testing it. The sparse empirical evidence on this area of research has put forth the role of exchange rate movements in generating nonlinearities. According to this literature, mainly, there are two potential sources of pass-through asymmetry. On one hand, asymmetry can arise from the *direction* of exchange rate changes i.e., in response to currency depreciations and appreciations. On the other hand, the extent of pass-through may also respond asymmetrically to the *magnitude* of exchange rate movements, i.e. depending on whether exchange rate changes are large or small. However, as pointed by MARAZZI et al. (2005), previous studies provide mixed results with no clear support for the existence of important nonlinearities. If the existing literature is not conclusive, there are two important caveats should be noted in this regard. First, ERPT is not depending exclusively on exchange rate changes, there are various factors, including macroeconomic variables, which might influence the pass-through mechanisms. Second, an appropriate econometric tool is required such as nonlinear regime-switching models where potential nonlinear behaviour in ERPT mechanism should be described correctly.

Consequently, in this chapter we propose to fill the gap in empirical evidence on the nonlinearities in ERPT. More precisely, we follow SHINTANI, TERADAHAGIWARA, and TOMOYOSHI (2009) and NOGUEIRA JR. and LEON-LEDESMA (2008) by using a STR models to estimate the extent of pass-through. We focus on “consumer-price pass-through”, i.e. the sensitivity of consumer prices to exchange rate changes. Unlike the cited studies, we are interested in the EA case since we expect that the different exchange rate arrangements experienced by the monetary union members would generate a nonlinear mechanism in ERPT. To our knowledge, there is no other

study has applied a nonlinear STR estimation approach in this context. The presence of nonlinearities is tested with respect to different macroeconomic determinants, namely the inflation environment, the direction and the size of exchange rate changes, the economic activity and the general macroeconomic stability.

According to our empirical results, we found that the degree of pass-through respond nonlinearly to the inflation environment, that is, ERPT is higher when the inflation level surpasses some limit. The time-varying ERPT coefficients point out that exchange rate pass-through has declined over time in the EA countries, this is due to the shift to a low-inflation environment. When considering the direction of exchange rate change as a potential source of nonlinearities - that is, whether ERPT asymmetrically to appreciation - we report mixed results with no clear evidence about the direction of asymmetry. This is not surprising since, in theory, an appreciation can lead to either a higher or lower rate of pass-through than depreciation. Next, we check the asymmetry of pass-through with respect exchange rate magnitude. We find that large exchange rate changes elicit greater pass-through than small ones. Results give a broad evidence of the presence of menu costs, when exchange rate changes exceed some threshold, firms are willing to pass currency movements through their prices. These findings seem to explain why ERPT was greater during the EMS Crisis and at the launch of the euro. Thereafter, the source of nonlinearities considered in our study is relative to business cycle. Our results provide a strong evidence of nonlinearity in 6 out of 12 EA countries with significant differences in the degree of ERPT between the periods of expansion and recession. However, we find no clear direction in this regime-dependence of pass-through to business cycle. In some countries, ERPT is higher during expansions than in recessions; however, in other countries, this result is reversed. Finally, we test whether periods of macroeconomic instability/confidence crisis may alter the extent of pass-through in a nonlinear way. In the light of the recent European sovereign debt crisis, we propose to use 10-year government bond yield differentials (versus Germany) as an indicator of macroeconomic instability. Our estimation is conducted only for the heavily-indebted EA economies i.e. the GIIPS group (Greece, Ireland, Italy, Portugal, and Spain). Results show that in periods of widening spreads, which corresponds to episodes of confidence crisis, the degree of ERPT is higher.

PART I

PASS-THROUGH TO IMPORT PRICES

Chapter 1

Measuring Exchange Rate Pass-Through to Import Prices in the Euro Area: An Update

1. Introduction

The study of the degree of Exchange Rate Pass-Through (ERPT) into import prices is of great policy interest in the euro area (EA) context. As import prices are a principal channel through which movements in the euro affect domestic prices and hence also the variability of inflation and output, the issue of pass-through has important implications for divergences in price level developments within the monetary union. A common exchange rate shock may impact differently on EA member states depending on their relative patterns of external exposure and openness to trade outside the euro zone. Thus, in achieving its target of medium-term price stability for the whole EA, the single monetary policy of the European Central Bank (ECB) must factor in the extent to which the euro exchange rate changes affect import prices.

Especially, the continuous depreciation of the euro (about 20 percent on a trade-weighted basis in the first two years) since its introduction has raised concerns that it might increase the risks to price stability. The weakening of the exchange rate of the

euro is likely to put upward pressure on import costs and producer prices, which can feed into higher consumer prices. The concern about the single currency depreciation affecting price stability has been clearly expressed by the monetary authority in the EA. In fact, the ECB cited the inflationary effects of a lower value of the euro as a factor behind its tightening of monetary policy in 2000.¹ This outcome has raised important questions regarding the magnitude and stability of ERPT since 1999, and, mainly, if EA members will be differentially affected by changes in the common external exchange rate. There has been a growing interest in the European ERPT in recent years. Studies conducted for the case of EA countries include HÜFNER and SCHRÖDER (2002), HAHN (2003); ANDERTON (2003), CAMPA, GOLDBERG, and GONZÁLEZ-MÍNGUEZ (2005), CAMPA and GONZÁLEZ (2006), FARUQEE (2006). A common drawback with these studies is the short time span available since the formation the euro in 1999. Thereby, in our study, we propose an update to ERPT elasticities using longer time period and more observations for the post-EA era.

Another important issue in the literature is related to the observed decline of the sensitivity of important prices to exchange rate movements in major of industrialized countries. Although the creation of the single currency euro area constituted a shift in both competition conditions and monetary policy, the European ERPT studies, including CAMPA, GOLDBERG, and GONZÁLEZ-MÍNGUEZ (2005) and CAMPA and GONZÁLEZ (2006), has failed to provide a strong evidence of reduction in pass-through. In fact, there are several factors which may lead to changing in the behaviour of ERPT, and, thus, would explain why the responsiveness of import prices has moved down markedly in the last two decades. An intriguing hypothesis was suggested by TAYLOR (2000) which explains that the shift towards more credible monetary policy and thus, a low-inflation regime would reduce the transmission of the exchange rate changes. This assumption is very appealing and has received strong empirical support in the recent literature (see e.g. GAGNON and IHRIG, 2004; BAILLIU and FUJII, 2004; CHOUDHRI and HAKURA, 2006). Nevertheless, the causes of the decline in pass-through are difficult to pin down with certainty, and there is an ongoing debate in this regard. In their sample of 23 OECD countries, CAMPA and GOLDBERG (2005) distinguish “micro-economic” from “macro-economic” explanations. The authors suggested that the product composition

¹See the statements given by the ECB in connection with Council monetary policy decisions between February and July 2000.

of a country's imports is by far more than macroeconomic factors such as inflation environment. That is, the shift in the composition of imports toward goods whose prices are less sensitive to exchange rate movements, such as differentiated manufactured products, is the most important driver of the marked fall of pass-through. Given the variability of the empirical findings, in this study we seek to shed light on these different issues by revisiting the euro zone case.

THE GOAL OF this chapter is to provide new up-to-date estimates of ERPT for 12 EA countries. First, we begin by estimating a static ERPT equation and analyze the main properties of the pass-through elasticities in our sample. This enables us to compare our results with those in the existing empirical literature on the EA, such as CAMPA, GOLDBERG, and GONZÁLEZ-MÍNGUEZ (2005), CAMPA and GONZÁLEZ (2006). These studies used few observations for the monetary union period (post-EA era), thus, their results are updated here. Following this individual estimates exercise, we assess the cross-country differences in our EA sample by investigating whether inflation level and degree of openness of an economy, as potential macro determinants, determine the magnitude of the pass-through. Next, as is typical in the empirical literature, we check the stability of the sensitivity of import prices to exchange rate movements over time. There are several reasons to believe that the degree of pass-through decreased since the inception of the euro in 1999. Among these explanations are the reduction of shares of imports exposed to exchange rate fluctuations and the increasing of the choice of the euro as a currency of denomination. Also, we have estimated our pass-through equation over different time period and compared results with those obtained over the benchmark period. Thereafter, we estimate a dynamic panel data model to provide an aggregate ERPT for the whole EA. The advantage of this framework is to allow us testing the influence of common events (experienced by the EA countries), such as the Exchange Rate Mechanism (ERM) crisis or the formation of the euro. Finally, we conduct a sectoral analysis in order to check for the importance of the composition of imports in determining the aggregate pass-through for a country.

To preview the results, over the estimation period of 1990-2010, we don't find a wide heterogeneity in the degree of pass-through across the 12 EA countries, in contrast to previous empirical works. This is not surprising since previous studies used too few observations for the EA era, while in our work, the time span for the analysis of the

post-EA era is rather long (until the end of 2010). Since the process of monetary union has entailed some convergence towards more stable macroeconomic conditions, it is expected to find a relative low and less dispersed ERPT across EA Member States. Concerning the macro determinants, we found a positive relationship between ERPT and inflation in line with TAYLOR (2000), while no significant role for the degree of openness, measured as the ratio of imports to GDP. Assessing the stability of pass-through elasticities, we find very weak evidence of decline around 1999, however, our results reveal that the fall in ERPT has started since the beginning of the 1990s. Within the dynamic panel data framework, we confirm the non-significant decline of the import-price sensitivity to exchange rate since the formation of the euro. However, the important role played by inflation environment was confirmed once again. We found that the responsiveness of prices to exchange rate fluctuations tend to decline in a low and more stable inflation environment. Moreover, our findings suggest that the weakness of the euro during the first three years of the monetary union has raised significantly the extent of pass-through. We pretend that this outcome would explain why the sensitivity of the import prices did not fall since 1999. Finally, using disaggregated import prices data, it appears that the product composition of imports would determine the aggregate ERPT of an economy, and thus, cross-country differences in pass-through rates are due to an heterogeneous industry composition of trade across countries.

The rest of the **Chapter 1** is structured as follows: Section 2 briefly reviews the literature on ERPT. In Section 3, a description of the behaviour of exchange rate and prices in the EA countries is given. Section 4 provides some theoretical considerations that underlie our empirical specification. Section 5 explains the empirical strategy and data sets used in this study. The results from our benchmark specification are presented in Section 6. In Section 7, we investigate for the potential decline in the pass-through elasticities. A dynamic GMM panel-data estimation is used in Section 8 to test for the influence of some macro factors on ERPT. In Section 9, a sectoral analysis of the degree of pass-through is provided. Section 10 concludes the chapter.

2. Overview of the literature

The mechanism of ERPT has long been of interest and has spawned many studies through the years. Acknowledging the large economic literature, we only survey some important studies which are cited frequently. The early literature was dominated by papers dealing with ERPT into import prices from a microeconomic perspective.² Along this vein, the industrial organization characteristics such as the presence of imperfect competition and price discrimination in international markets are main factors explaining the incomplete pass-through. In seminal papers, DORNBUSCH (1987) and KRUGMAN (1987) justified incomplete pass-through as arising from firms that operate in a market characterized by imperfect competition and adjust their markup in response to an exchange rate shock. As is well-known, the markup depends on the elasticity of demand for a given product, which is, in turn, determined by competitors' prices. Facing a change in the exchange rate, producers can decide whether and to what degree the markup should absorb these changes. When the currency of the importing country is depreciating, a foreign firm might cut its price by reducing its markup in order to stabilize its price in terms of the importing country's currency, then pass-through is less than complete.³

Although the degree of pass-through has played a central role in debates in international economics for a long time, the question of whether pass-through can be influenced by the macroeconomic environment and in particular the role of monetary policy, is a more recent occurrence. The emerging macro literature has focused on the issue of the relatively widespread and on-going decline in ERPT. A popular view in this regard has put forward in particular by TAYLOR (2000). The author provides a model where lower pass-through is caused by lower perceived persistence of inflation.

²It is noteworthy that most of the early pass-through literature has focused on traded goods prices such as import or export prices and very few on consumer-price ERPT

³It is important to note that the micro-based literature has a partial-equilibrium approach, that is, they focus on the response of prices to an exogenous movement in the nominal exchange rate. As an alternative to this approach, structural vector autoregressions (VAR) have become increasingly popular as a method to estimate the exchange rate pass-through (see e.g. MCCARTHY, 2007). A motivation for using the structural VAR approach is that it takes explicit account of the endogeneity of the exchange rate and permits the estimation of pass-through to a set of prices, such as import prices, producer prices and consumer prices, simultaneously.

The more persistent inflation is, the less exchange rate movements are perceived to be transitory and the more firms might respond via price-adjustments. Thus, countries with credible and anti-inflationary monetary policies tend to experience lower ERPT.⁴ Several empirical studies were very supportive of Taylor's view.⁵ For instance, GAGNON and IHRIG (2004) explore the relationship between pass-through to consumer prices and inflation stabilization in a sample of 20 industrialized countries over the period 1972-2000. They find that the pass-through generally declined in the 1990s and that countries with low and stable inflation rates tend to have low estimated rates of pass-through.

Besides, Taylor's hypothesis has been theoretically examined in the context of the new open-economy macroeconomics.⁶ In this type of framework, ERPT will depend on different pricing strategies, i.e. whether the foreign exporter follows a producer currency pricing (PCP) or local currency pricing (LCP) strategy. When prices are determined in the exporter's currency (PCP), pass-through tends to be much larger than when prices are set in the importer's currency (LCP). In the extreme case of a purely exogenous exchange rate shock, exchange rate pass-through would be one under producer currency pricing and zero under local currency pricing. It is worth noting that this literature provide a reconciliation between macroeconomic and microeconomic factors. In this vein, DEVEREUX, ENGEL, and TILLE (2003) developed a dynamic general-equilibrium model linking the extent of pass-through to monetary policy. They conclude that countries with low relative exchange rate variability and relatively stable monetary policies would have their currencies chosen for transaction invoicing. In this case, prices are sticky in the currency of the importing country (local currency pricing (LCP)), and pass-through tend to be low. However, exchange rate pass-through would be higher for importing countries with more volatile monetary policy. Prices will be preset in the currency of the exporter, i.e. prevalence of producer currency pricing strategy, and then ERPT will tend to be high.⁷ However, IHRIG, MARAZZI, and ROTHENBERG (2006) caution against the local currency pricing hypothesis. They argued that exporters

⁴This explanation seems to bear more on pass-through to consumer prices than on pass-through to import prices.

⁵Most of these studies consider the pass-through to consumer prices.

⁶This strand of literature is based mainly on the seminal Redux model OBSTFELD and ROGOFF (1995) incorporating imperfect competition and price inertia into a dynamic general-equilibrium open-economy model.

⁷Same finding was provided by DEVEREUX and ENGEL (2002).

may choose to invoice in the currency of the destination market to shield the price paid by its clients from exchange-rate movements in the medium-term. However, over the long run, in the face of a protracted appreciation of the exporter's currency, it will have to adjust its local currency price to keep its margins from turning negative.

However, there is a serious debate on the prevalence of macroeconomic factors vs. microeconomic factors. GOLDBERG and TILLE (2008) provide empirical evidence suggesting that the choice of invoicing currency is influenced more by the product composition of trade than by macroeconomic factors. If trade is largely in homogeneous, the role of macroeconomic variability in invoice currency choice is substantially damped. For producers, the most important driver of invoice currency selection will be the need to have their goods priced in the same way as other competing producers price their products. The same view was emphasized by CAMPA and GOLDBERG (2005) in their studies of import-price pass-through in 23 OECD countries. According to the authors, the macroeconomic variables - levels of inflation, money growth rates or country size - are weakly correlated with changes in pass-through, and hence they are not of first order importance in explaining pass-through evolution within the OECD over the past 25 years. Whereas, they found a strong evidence that shift in the composition of imports toward goods whose prices are less sensitive to exchange rate movements has contributed to a fall in the pass-through in many countries in the 1990s. MARAZZI et al. (2005) take a somewhat different view. According to the authors, "the Campa-Goldberg compositional-change hypothesis" may explain some, but certainly not the lion's share, of the decline in pass-through in the United States. This phenomenon can only explain about one-third of the decline in pass-through to U.S. import prices. MARAZZI et al. (2005) also suggest that China's surging exports to the U.S. may be partly responsible for the low levels of observed pass-through in US economy.

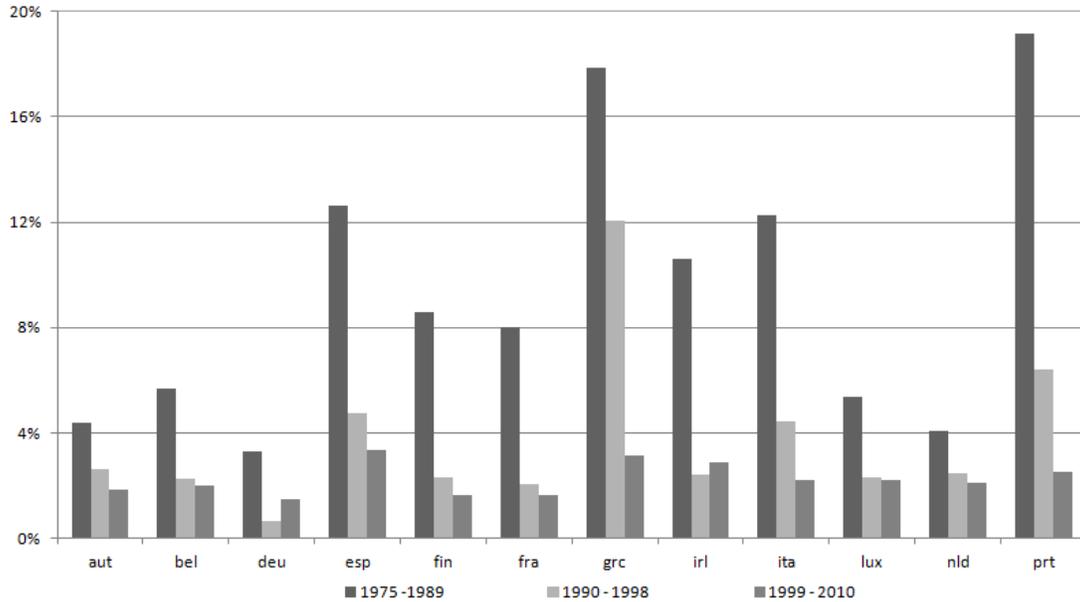
A host of other hypotheses have also been put forward as factors causing incomplete or declining ERPT to import prices. MANN (1986) documented that the increased usage of exchange-rate hedges may shield a firm from exchange rate shocks allowing them to avoid passing such shocks to consumers. Although hedging can allow firms to postpone passing through an exchange rate shock, but in the long-run a sufficiently large and permanent exchange rate shock will have to be passed through to importers. Another argument for incomplete pass-through was articulated by BODNAR,

DUMAS, and MARSTON (2002) and is related to cross-border production arrangements. If production takes place in several stages across many countries, then the costs of producing the final good are incurred in several currencies. This can explain incomplete pass-through as long as all of these currencies do not experience a common appreciation against the export destination's currency. We finally note a recent paper by GUST, LEDUC, and VIGFUSSON (2010) that proposes the process of international globalization itself may induce a fall in pass-through. In their model, lower trade costs (interpreted broadly as increased globalization) increase exporting firm's relative markups which in turn allow their prices to be less sensitive to exchange rates yielding lower pass-through.

3. Exchange rate and prices dynamics in the EA countries

Before estimating the degree of ERPT across the 12 EA countries, in this section we try to shed some light on the behavior of the key macro variables in the European context. Beginning with the inflation levels as displayed in Figure 1.1 throughout three sub-sample periods: 1975-1989, 1990-1998 and 1999-2010. We note a very high inflation rates on average in the 1970s and the 1980s, especially, in the so-called "peripheral " EA countries, i.e. Greece, Ireland, Italy, Portugal and Spain, with a double-digit level of inflation.⁸ During the 1990s and 2000s, we note a steady decline in the inflation environment in all of the monetary union members. Greece still have more than one-digit level of inflation between 1990-1998, i.e. during the first and the second stage of European Monetary Union (EMU), but the rate has fallen markedly since the inception of the single currency. In the other extreme, we point out that Germany has had a stable consumer prices changes all along the three sub-sample periods. The trend towards low and stable inflation, that we have in Figure 1.1, is a result of the process of inflation convergence that was started since the implementation of Maastricht treaty. Knowing the strong relationship between ERPT and inflation environment, as stated by TAYLOR (2000), this would have a significant impact of the degree of transmission of exchange rate given the various inflation history of our country sample.

⁸Recently, in the context of the European sovereign-debt crisis, the term PIIGS was employed to label these heavily-indebted economies.

Figure 1.1: Mean rate of inflation in the EA countries

Note: Inflation is computed as the quarterly year-on-year changes of consumer prices. Data are from International Financial Statistics of IMF.

Next, we inspect the behavior of the exchange rate changes (Δe), the import prices inflation (Δpm) and consumer prices inflation (Δcpi). For these macroeconomic variables, the quarterly average for the mean and for the standard deviation over 1990-2010 are summarized in Table 1.1. The exchange rate is defined in terms of local currency units per unit of the foreign currency, thus, an increase corresponds to a depreciation. Average import-price inflation were the largest in Belgium followed by Greece and Italy, while in terms of volatility, i.e. standard deviation, Greece and Ireland have the highest value. As regards consumer prices, we find the same pattern as in Figure 1.1: peripheral countries such as Spain, Greece, Ireland and Portugal has the less stable inflation rates both in value and volatility terms. However, the so-called “core” EA countries, such as Austria, Belgium, Germany and France exhibit low average inflation rates and, consequently, a more stable macroeconomic conditions than peripheral EA members.

Thereafter, we have explored the link between inflation and exchange rate variation. In Table 1.1, we provide an overview of the simple correlation between

quarterly changes in inflation and nominal exchange rates over 1990-2010. Regarding import prices inflation, the expected positive correlation with the currency movements is evident for all of the EA countries with the exception of Austria. The tightest relationships are found in Spain, Ireland and Italy, while the level of the correlation coefficient is notably low in the case of France. For Belgium, although the high average import-price inflation over the sample period, the relationship with exchange rate is weaker than expected.

Table 1.1: Summary statistics over 1990-2010

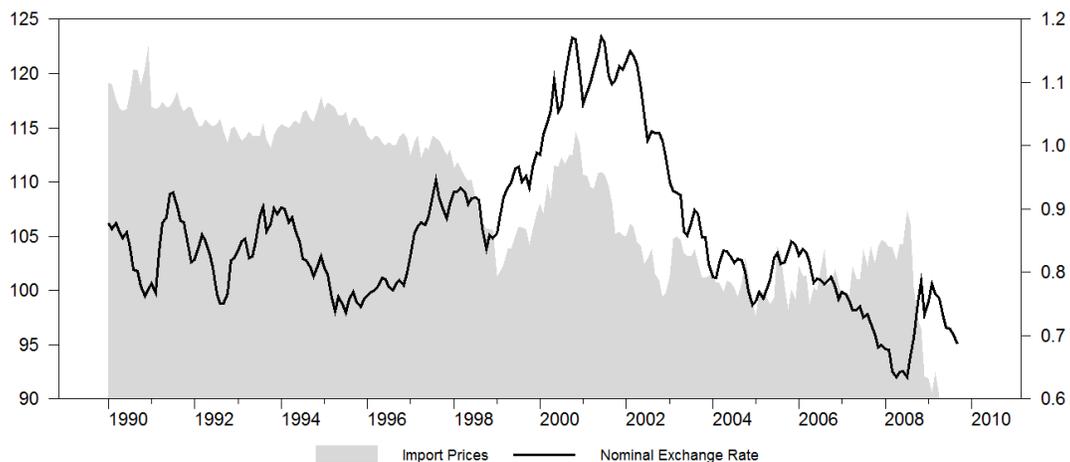
Country	Correlation ($\Delta pm, \Delta e$)	Correlation ($\Delta cpi, \Delta e$)	Mean (%)			Standard Deviation (%)		
			Δpm	Δcpi	Δe	Δpm	Δcpi	Δe
Austria	-0,07	-0,20	0,19	0,54	-0,33	0,62	0,57	0,88
Belgium	0,27	-0,06	1,56	0,52	-0,32	1,57	0,47	1,27
Germany	0,36	-0,02	-0,02	0,45	-0,18	1,22	0,47	1,45
Spain	0,62	0,11	0,42	0,84	0,21	2,08	0,80	1,67
Finland	0,33	0,18	0,36	0,45	0,27	1,97	0,44	2,39
France	0,13	-0,06	-0,01	0,44	-0,09	1,34	0,37	1,20
Greece	0,41	0,59	-0,82	1,61	0,40	5,73	1,94	1,68
Ireland	0,62	0,04	-0,38	0,62	0,06	3,89	0,79	2,19
Italy	0,65	0,26	0,76	0,76	0,34	2,03	0,46	2,33
Luxembourg	0,20	-0,06	0,71	0,55	-0,09	2,12	0,50	1,12
Netherlands	0,26	-0,09	0,10	0,55	-0,06	1,28	0,43	1,41
Portugal	0,28	0,22	0,33	0,98	0,05	1,49	0,96	1,23

Source: OECD & personal calculation.

With regard to consumer-price inflation, the co-movement with the nominal exchange rate is much lower than recorded with import-price inflation. In the half of the EA countries, we have a wrong (negative) sign of correlation coefficients. However, on the other side, we found that exchange rate depreciation is positively associated with higher inflation rates of consumer prices, especially, for Greece, Italy and Portugal. On the basis of the inflationary record, one would expect this positive relationship. Nevertheless, these results represent a statistical correlation without specific economic interpretation in terms of ERPT. Therefore, in section 5, we provide an econometric analysis using more economically meaningful specifications to assess the relationship between exchange rates and prices.

Another point that deserves more attention is related to the sizable swings of the euro since 1999. The single currency depreciated strongly against the U.S. dollar in 1998 through 2001, followed by an appreciation of roughly the same magnitude between 2002 and 2004. During the first two years of his existence, the euro depreciated by approximately 25% in a nominal trade-weighted basis, and since the second quarter of 2002, it started to appreciate regaining about 20% of its value by the end of 2004. Now, an important question is how and to what extent these large movements in exchange rates are reflected in prices.

Figure 1.2: Import prices and the nominal exchange rate (against US dollar) in France



Source: International Financial Statistics of IMF.

As shown in Figure 1.2 which tracks both nominal exchange rate and the import prices in France (over 1990-2010), these dramatic changes in the value of the euro seem to induce a considerable increasing in the import prices since 1999. It is expected that such movements can put substantial pressures on foreign producers to adjust their prices accordingly. The concern about the exchange rate affecting price stability during this episode has been clearly expressed in the European Central Bank (ECB) reactions. The contractionary monetary policy in 2000 was a response to the inflationary effects of the weakness of the euro, while the loosening in 2003 is due to the disinflationary effects of a strong euro. Otherwise, it should be noted that some industrialized countries has also experienced a considerable depreciation of the exchange rate without domestic prices

being affected as much as expected. This was the case of Canada, Sweden, and the United Kingdom in the 1990s. Therefore, to ensure the potential strong relationship between exchange rate and import prices during the first years of the creation of the EA, a relevant econometric methodology must be employed before drawing any definite conclusions.

4. Theoretical Framework

The theoretical framework used here follows FEENSTRA (1989) and COUGHLIN and POLLARD (2004). The model is set in the context of a price-discriminating monopolist and it is a partial equilibrium. Considering a domestic importing country that imports a differentiated good q^m from a monopolist foreign firm which is facing competition from a substitute good z in the importing country. Assuming that the differentiated product q^m is weakly separable from other goods in the consumer's utility function, import demand of good q^m can be expressed as follow: $q^m(p^m, p^z, Y^m)$, where p^m denotes the import price of q^m in the domestic currency, p^z is the domestic currency price of z and Y^m is the income or expenditures on all goods in the importing country. At the same time, foreign exporter firm produces good q^x for sale in his local market with the following local (foreign) demand: $q^x(p^x, Y^x)$, where p^x is the foreign currency price of the good and Y^x is the income or expenditures on all goods in Foreign.

In this economy, the good q is produced only in the foreign country and inputs are allowed to come from both domestic and foreign countries. Thus, factor prices in the foreign country, w^* , will depend on the exchange rate, e (number of units of importing country's currency per unit of foreign currency). The foreign firm's cost function is given by $c(Q, w^*(e))$, where Q is the total quantity produced for both domestic and foreign markets ($Q = q^m + q^x$). Costs are assumed to be homogeneous of degree one in factor prices, so they can be written as $c(Q, w^*(e)) = w^*(e)\phi(Q)$. The foreign firm maximizes profits in its own currency, treating z and Y^m as exogenous.⁹

⁹Foreign and domestic firms are assumed to act as Bertrand competitors.

Then, the profit maximization problem can be stated as:

$$\max_{p^x, p^m} \Pi = p^x q^x + e^{-1} p^m q^m - w^*(e) \phi(Q) \quad (1.1)$$

The first-order condition for (1.1) is:

$$p^x : q^x + p^x \frac{\delta q^x}{\delta p^x} - w^* \phi' \frac{\delta q^x}{\delta p^x} = 0 \quad (1.2)$$

$$p^m : e^{-1} q^m + e^{-1} p^m \frac{\delta q^m}{\delta p^m} - w^* \phi' \frac{\delta q^m}{\delta p^m} = 0 \quad (1.3)$$

Equations (1.2) and (1.3) can be rewritten as:

$$p^x : \frac{\delta q^x}{\delta p^x} \left[p^x \left(1 - \frac{1}{\varepsilon^x} \right) - w^* \phi' \right] = 0 \quad (1.4)$$

$$p^m : \frac{\delta q^m}{\delta p^m} \left[e^{-1} p^m \left(1 - \frac{1}{\varepsilon^m} \right) - w^* \phi' \right] = 0 \quad (1.5)$$

where $\varepsilon^i = - \left(\frac{\delta q^i p^i}{\delta p^i q^i} \right)$ is the elasticity of demand with respect to price for $i = x, m$.

Knowing that markup over marginal cost is defined as $\mu^i = \left(\frac{\varepsilon^i}{\varepsilon^i - 1} \right)$, the first order

conditions become:

$$p^x : \frac{\delta q^x}{\delta p^x} \left[\frac{p^x}{\mu^x} - w^* \phi' \right] = 0 \quad (1.6)$$

$$p^m : \frac{\delta q^m}{\delta p^m} \left[\frac{e^{-1} p^m}{\mu^m} - w^* \phi' \right] = 0 \quad (1.7)$$

Therefore:

$$p^x = w^* \phi' \cdot \mu^x \quad (1.8)$$

$$p^m = e \cdot w^* \phi' \cdot \mu^m \quad (1.9)$$

We see that solving profit maximization yields the standard condition that the price in each market, i.e. foreign and domestic, is determined by a market specific markup, μ^i , over common marginal cost, $w^* \phi'$. Our primary focus is on the equation (1.9). This latter shows that the import price p^m (which is expressed in the importing country's currency) depends on three factors: the bilateral exchange rate between importer and exporter, the marginal cost and the markup of price over marginal cost. We note that the exporter's marginal cost and markup may change independently of the exchange rate. For instance, a change in the cost of a locally provided input (in the foreign country) can shift the marginal cost. Also, adjustments in markups may occur in response to changes in variables specific to importing country, mainly, demand conditions Y^m and the price of the competing product p^z , so that: $\mu^m = \mu^m(Y^m, p^z)$.

Supposing that marginal costs are constant, $w^* \phi'' = 0$, we can derive the ERPT elasticity as follow:¹⁰

$$\text{ERPT} = \frac{\delta p^m}{\delta e} \frac{e}{p^m} = \frac{1 + \eta^{w^*e}}{1 - \eta^{\mu m}} \geq 0 \quad (1.10)$$

where $\eta^{w^*e} = \frac{\delta w^*}{\delta e} \frac{e}{w^*} \leq 0$ and $\eta^{\mu m} = \frac{\delta \mu^m}{\delta p^m} \frac{p^m}{\mu^m} \leq 0$ are the elasticity of factor prices with respect to the exchange rate and the elasticity of the markup with respect to the price in domestic country currency, respectively.

According to equation (1.10), we see that pass-through elasticity depends crucially on the behavior of the marginal cost and markup. In general, ERPT is positive in the sense that a depreciation in the importing country's currency ($\uparrow e$) increases the import price of good; while an appreciation of the currency value ($\downarrow e$) raises the price of imported good.¹¹ Equation (1.10) suggests that full pass-through (ERPT=1) is a special case. If marginal cost is not affected by the exchange rate fluctuations ($\eta^{w^*e} = 0$), i.e. foreign producer uses only local inputs in the production process, and the markup is constant ($\eta^{\mu m} = 0$), pass-through would be complete. In the case of higher sensibility of marginal costs to exchange rate, that is when $\eta^{w^*e} = -1$, there will be zero ERPT. Also, in the case of extreme sensibility of markup to domestic currency import price ($\eta^{\mu m} \rightarrow -\infty$), foreign exporter offset exchange rate changes by adjusting markup, and then ERPT tend to be zero.

5. Empirical framework and data

In this section, we present the empirical model used to estimate the degree of pass-through which stems from the analytical framework presented above. As stated by the import price equation (1.9), in estimating ERPT it is necessary to isolate the

¹⁰The derivations of ERPT elasticity is given in the appendix A.1 in more details.

¹¹As explained by COUGHLIN and POLLARD (2004), this can be generalized as long as marginal costs are nondecreasing in output, $\phi'' \geq 0$. However, in the case of decreasing marginal costs, $\phi'' < 0$ and the elasticity of input costs with respect to the exchange rate is $\eta^{w^*e} < -1$, then ERPT may be negative.

exchange rate effect from other effects, i.e. the exporter's cost shifter, importer's demand conditions, and the price of the domestic competitor. Thus, we can capture the arguments of the import price equation (1.9) through a log-linear regression specification similar to that tested throughout the ERPT literature:

$$p_t^m = \beta_0 + \beta_1 e_t + \beta_2 w_t^* + \beta_3 Z_t + \varepsilon_t, \quad (1.11)$$

where p_t^m are domestic currency import prices, e_t is the exchange rate, w_t^* variable representing exporter costs, and Z_t is a vector including demand conditions and competitors prices in the importing country with other controls. As discussed by CAMPA and GOLDBERG (2002), biased estimates of the pass-through coefficient could arise if foreign costs or proxies for markup are correlated with exchange rates but omitted from the regression. Variants of equation (1.11) are widely used as empirical specifications in the pass-through literature.¹² While the general approach of is very similar in the pass-through studies, there are a few differences between them regarding the specification and the list of control variables. Our primary concern in this study is the pass-through elasticity which corresponds to the coefficient on the exchange rate β_1 in the empirical model just outlined. The parameter β_1 is expected to be bounded between 0 and 1. A one-for-one pass through to changes in import prices, known as a complete ERPT, is given by $\beta_1 = 1$. In this case, exporters let the domestic currency import prices affected by exchange rate move. While, when exporters adjust their markup, a partial or incomplete ERPT occurs and $\beta_1 < 1$.

In our empirical work, the degree of pass-through into import prices is estimated for 12 EA countries: Austria, Belgium, Spain, Germany, Finland, France, Greece, Ireland, Italy, Luxembourg, Netherlands, and Portugal. We have the same country sample as in CAMPA, GOLDBERG, and GONZÁLEZ-MÍNGUEZ (2005) and CAMPA and GONZÁLEZ (2006). However, for the latter studies, time period estimation covers only up until the mid-2004. In our analysis, we provide an *up-to-date* of ERPT estimates for the main members of the monetary union. The period of estimation corresponds to the interval that spans from 1990:3 to 2010:4 using quarterly data. This allows us to

¹²See GOLDBERG and KNETTER (1997) for a survey of this literature.

compare our estimates with existing results for EA countries. For each country, data was collected following a cascade order and choosing when possible only one institutional source, namely IMF's *International Financial Statistics* and OECD's *Main Economic Indicators* and *Economic Outlook*, in that order.

Concerning our dependent variable, i.e. the domestic import prices, we use the price of non-commodity imports of goods and services. This represents import prices of core goods by excluding primary raw commodities because of their marked volatility. For all countries for the exchange rate we employed the nominal effective trade weighted series, with an increase means depreciation of the national currency, and a decrease means appreciation. Next, the marginal costs of foreign producers are difficult to measure since they are not directly observable, and thus need to be proxied. A conventional practice is to use a weighted average of trade partners' costs as in CAMPA and GOLDBERG (2005) and BAILLIU and FUJII (2004). Following this, the foreign costs of each EA country's major trade partners is derived implicitly from the nominal and real effective exchange rate series as follows: $w^* \equiv q_t - e_t + ulc_t$, where ulc_t is the domestic unit labour cost (ULC) and q_t is the ULC-based real effective exchange rate. Given that the nominal and real effective exchange rate series are trade weighted, this proxy provides a measure of trading partner costs, with each partner weighted by its importance in the importing country's trade. As regards foreign firm's markup, in our benchmark specification, we use the output gap, as the difference between actual and HP-filtered gross domestic product (GDP), to proxy for changes in domestic demand conditions. To check the robustness of the benchmark model, in addition to the output gap, we have included the domestic producer prices ppi_t as a proxy for the competitors prices in the importing country (similar to OLIVEI (2002) and BUSSIÈRE (2012), among others). Also, to check the reliability of the output gap as a good proxy for the domestic conditions, the real GDP (as in CAMPA, GOLDBERG, and GONZÁLEZ-MÍNGUEZ (2005)) can be used instead. Furthermore, as is well-known, changes in the exchange rate also influence import prices indirectly through their effects on commodity prices. To consider such channel, as robustness test, we can include oil prices oil_t (in dollar US) as an additional explanatory variable in the pass-through equation. As explained by IHRIG, MARAZZI, and ROTHENBERG (2006), when it was not possible to find import prices of core goods that exclude all primary raw commodities, the inclusion

of commodity prices indexes, such as oil prices, as independent variables should mitigate some of the noise generated by these volatile components. All the robustness tests with different specifications of ERPT equation are reported in Appendix A.3.¹³

Another concern in the ERPT equation is related to the fact that foreign costs and the exchange rate would have the same coefficient, i.e. $\beta_1 = \beta_2$, as predicted by the theoretical framework in HOOPER and MANN (1989). In practice, this restriction does not necessarily hold, since that exchange rates are more variable than costs, and thus, the extent to which they are passed on prices may differ (see ATHUKOROLA and MENON (1995) for a discussion). To test the restriction whether parameters at the exchange rate and foreign costs are equal, Wald test will be conducted subsequently.

Finally, we check the stationarity of our key variables. Augmented Dickey Fuller (ADF) and ZIVOT and ANDREWS (1992) stationary tests in Table A.1 in Appendix A.2 indicate that most of the variables are integrated of order one $I(1)$, except the output gap which is by construction a stationary variable.¹⁴ Given that the time series properties of the data, i.e. non-stationarity, we investigate the possibility cointegration between variables in levels. To achieve this, in addition to ENGLE and GRANGER (1987) test (EG hereafter), we also employ GREGORY and HANSEN (1996) test (henceforth GH) which allows for structural breaks in the cointegrating vector. As reported in Table A.2 in Appendix A.1, overall, there is a weak evidence of possible long-run equilibrium relationships among the variables; the residuals of ERPT equation in levels are non-stationary for most of the countries in our sample. This is not surprising, given that most researchers have not found evidence in favour of cointegrating relations between the variables (see *inter alia* CAMPA and GOLDBERG (2005), CAMPA, GOLDBERG, and GONZÁLEZ-MÍNGUEZ (2005) and CAMPA and GONZÁLEZ (2006)). Therefore, we have taken for estimation the first differences of the variables in order to control for the possibility of non-stationarity in the time series variables. Also, because the data are not seasonally adjusted, quarterly dummy variables are included to capture any seasonal

¹³The additional controls, i.e. producer prices and oil prices, are not considered in our benchmark model in order to avoid multicollinearity issues. For instance, we found that the correlation between the output gap measure and the producer prices is quite high.

¹⁴ZIVOT and ANDREWS (1992) test allow for one single break under the alternative hypothesis.

effects. The first-difference version of equation (1.11) has the following form:

$$\Delta p_t^m = \beta_0 + \beta_1 \Delta e_t + \beta_2 \Delta w_t^* + \beta_3 gap_t + \text{quarterly dummies} + \varepsilon_t, \quad (1.12)$$

This ends up estimating an import-price inflation equation. The estimation methodology applied on (1.12) is ordinary least squares (OLS).¹⁵

6. ERPT results from the benchmark model

In this section, we provide the estimation results from equation (1.12) over 1990–2010 as summarized in Table 1.2.¹⁶ Overall, the estimation results show that the coefficients of the key variables are statistically significant with expected signs, namely exchange rate depreciation and foreign costs affect positively the domestic currency import prices. The exception is the output gap which is found to be positively significant only for 4 out of 12 EA countries.¹⁷ This puzzling result was pointed out throughout the ERPT literature (see e.g. BUSSIÈRE, 2012). We think that a more thorough econometric analysis would improve the results. As explained by MENON (1995) in his exhausting discussion of ERPT literature, most of the empirical studies employ an OLS estimation technique which does not properly take into account the time series properties, namely the non-stationarity of the data. In the Chapter 2, we provide a panel cointegration approach which enables us to restore the long-run equilibrium relationship, and at the same time gives us more significant coefficients on the domestic demand.

¹⁵When include producer prices in equation (1.12), the use of instrumental variable estimator would be more accurate. In fact, the domestic firms compete against the exporting firm taking the level of import prices into account, thus producer domestic prices need to be treated as an endogenous regressor (see BUSSIÈRE (2012)). The instrumental variable technique using lagged domestic product prices as instruments shows that the results are very similar to OLS estimator.

¹⁶Because of data availability, the estimation period is 1990:3-2010:3 for Austria and Ireland, and 1990:3-2010:2 for Greece.

¹⁷Higher domestic demand would tend to raise import prices.

Table 1.2: Estimation results from pass-through equation

	Austria	Belgium	Germany	Spain	Finland	France
Constant	0,028 (0,000)	-0,001 (0,808)	-0,004 (0,026)	0,000 (0,885)	-0,006 (0,126)	-0,004 (0,018)
Δe_t	0,287 (0,000) [0,000]	0,428 (0,000) [0,000]	0,379 (0,000) [0,000]	0,553 (0,000) [0,000]	0,323 (0,000) [0,000]	0,372 (0,000) [0,000]
Δw_t^*	0,428 (0,000)	0,607 (0,000)	0,583 (0,000)	0,664 (0,000)	0,515 (0,000)	0,624 (0,000)
gap_t	-0,014 (0,888)	0,311 (0,003)	0,024 (0,468)	0,100 (0,359)	0,039 (0,622)	0,061 (0,480)
Observations	81	82	82	82	82	82
R^2	0,891	0,572	0,703	0,590	0,320	0,653
Wald test	10,363	19,308	35,429	5,647	10,338	72,496
p -value	(0,002)	(0,000)	(0,000)	(0,020)	(0,002)	(0,000)
	Greece	Ireland	Italy	Luxembourg	Netherlands	Portugal
Constant	0,009 (0,001)	0,006 (0,114)	0,005 (0,026)	0,011 (0,010)	-0,001 (0,737)	0,002 (0,467)
Δe_t	0,476 (0,000) [0,000]	0,423 (0,000) [0,000]	0,586 (0,000) [0,000]	0,448 (0,000) [0,000]	0,404 (0,000) [0,000]	0,460 (0,000) [0,000]
Δw_t^*	0,721 (0,000)	0,329 (0,002)	0,771 (0,000)	0,656 (0,000)	0,637 (0,000)	0,693 (0,000)
gap_t	0,090 (0,266)	0,145 (0,054)	0,170 (0,048)	-0,088 (0,319)	0,119 (0,032)	0,019 (0,776)
Observations	80	81	82	82	82	82
R^2	0,607	0,422	0,795	0,292	0,734	0,649
Wald test	42,168	2,783	29,763	9,756	94,667	50,018
p -value	(0,000)	(0,099)	(0,000)	(0,003)	(0,000)	(0,000)

Note: Estimation are based on equation 1.12. Numbers in parentheses are p -values. For the exchange rate coefficient, p -values in parentheses are based on the null hypothesis of zero ERPT, i.e. $H_0 : \beta_1 = 0$, while p -values in square brackets corresponds to the null of full ERPT, i.e. $H_0 : \beta_1 = 1$. Wald test is performed for $H_0 : \beta_1 - \beta_2 = 0$.

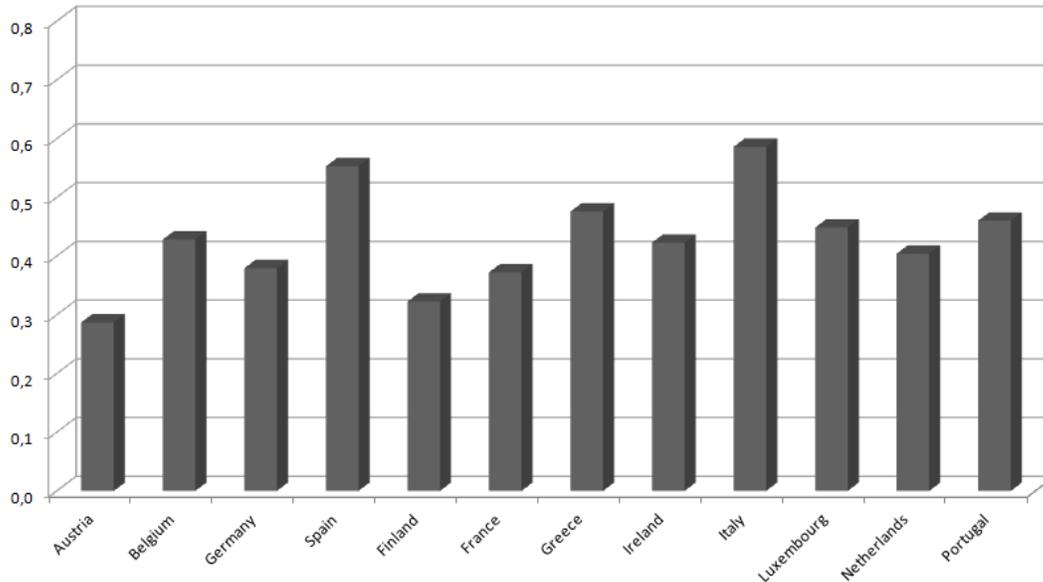
Turning to the estimated ERPT coefficients, we note that ERPT elasticities are positively significant in all EA countries and bounded between 0.28% (for Austria) and 0.59% (for Italy). Contrary to previous empirical studies, we don't find a wide heterogeneity in the degree of pass-through across the 12 EA countries (see Figure 1.3). For instance, a significant degree of variability in ERPT estimates across EA countries was reported in CAMPA, GOLDBERG, and GONZÁLEZ-MÍNGUEZ (2005) and CAMPA and GONZÁLEZ (2006). Also, in our study we find that the average of the exchange rate transmission into the aggregate import prices is equal to 0.43%. In other words, a one percent increase in the rate of depreciation of domestic currency raise the import prices

by 0.43 percent in average in our EA sample. Our estimates of ERPT are slightly lower in comparison with CAMPA, GOLDBERG, and GONZÁLEZ-MÍNGUEZ (2005) and CAMPA and GONZÁLEZ (2006). In the latter papers, the short-run pass-through elasticities are close to 0.66% in average for 11 EA countries.¹⁸ This outcome is not surprising since the mentioned studies used too few observations for the EMU era (CAMPA, GOLDBERG, and GONZÁLEZ-MÍNGUEZ (2005) until mid-2004 and CAMPA and GONZÁLEZ (2006) until the end of 2001). However, in our work, the time span for the analysis of the post-EA era is rather long (until the end of 2010). Since the process of monetary union has entailed some convergence towards more stable macroeconomic conditions, it is expected to find a relative low and less dispersed ERPT across EA Member States.

Otherwise, we can test for the prevalence of local currency pricing (LCP) versus producer currency pricing (PCP) strategy. LCP represents a null hypothesis of zero pass-through, i.e. $H_0 : \beta_1 = 0$, whereas PCP implies a pass-through of unity, i.e. $H_0 : \beta_1 = 1$. Our results show that both LCP and PCP hypotheses are strongly rejected in all EA countries. According to our results, partial ERPT is the best description for import prices responsiveness to exchange rate changes in our country sample. For 23 OECD countries, CAMPA, GOLDBERG, and GONZÁLEZ-MÍNGUEZ (2005) support this view in the short-run; import prices reaction are significantly different from zero in 20 out of 23 countries and significantly different from one for 18 out of 23 countries. However, the authors found that LCP hypothesis is not rejected for Austria, Belgium and Ireland, while the hypothesis of full ERPT (PCP strategy) is accepted for Finland. Nevertheless, the time span in CAMPA, GOLDBERG, and GONZÁLEZ-MÍNGUEZ (2005) covers the period from 1975 through 2003, which contains a longer period prior to EMU but less observations in the post-EA period.¹⁹

¹⁸In CAMPA, GOLDBERG, and GONZÁLEZ-MÍNGUEZ (2005) and CAMPA and GONZÁLEZ (2006), Belgium and Luxembourg are treated as a single country.

¹⁹As predicted by some theoretical models, we have tested for the restriction of equality on coefficients of exchange rate and foreign prices, i.e. $H_0 : \beta_1 = \beta_2$ (see HOOPER and MANN, 1989). According to Wald test results (last lines in Table 1.2), the hypothesis of equal parameters is rejected for our entire country sample. This outcome is in line with most of empirical studies which argue that exchange rates are more volatile than costs, and, thus, imposing such restriction does not necessarily hold (see ATHUKOROLA and MENON, 1995).

Figure 1.3: ERPT elasticities in EA countries over 1990:3-2010:4

Source: Personal calculation.

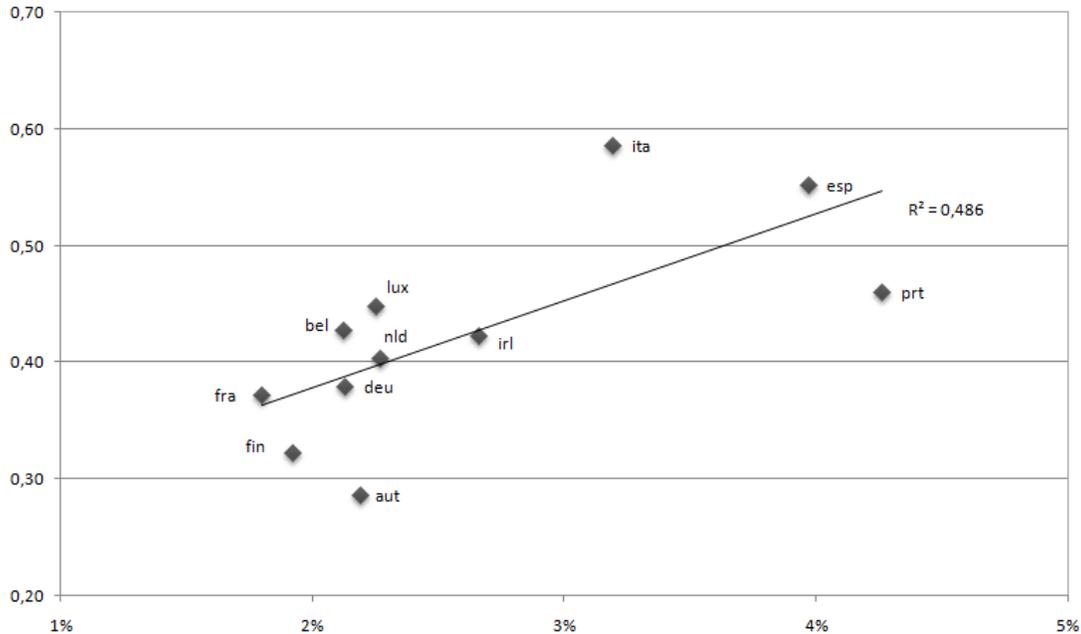
Regarding the robustness checks, of the results obtained from the equation (1.12) seems to be robust to the inclusion of producer prices as an additional explanatory variables to proxy for competitors prices in the importing country (see Table A.3 in Appendix A.3). We point out that the coefficients of the variable are quite lower and not significantly different from zero in all case except for Greece. Similarly, when we introduce oil prices in the regression, this does not alter significantly the results of the benchmark specification (see Table A.4 in Appendix A.3). However, we underline that ERPT coefficients are slightly lower when oil prices is introduced in equation (1.12). This is not surprising since when commodity prices such as oil prices are excluded from the regression, the pass-through coefficients capture both the direct effect of the exchange rate on import prices and the indirect effect operating through changes in commodity prices. Thus, taking into account this latter channel would slightly lower the ERPT elasticity (see MARAZZI et al., 2005; IHRIG, MARAZZI, and ROTHENBERG, 2006; MARAZZI and SHEETS, 2007, for a discussion). Finally, we have replaced the output gap by the rate of growth of real GDP in (1.12) as in CAMPA and GOLDBERG (2005). The results are still the same, i.e. the coefficients on output growth are

insignificant in most cases, and even when it is, it does not affect the other coefficients. Broadly, we can say that our benchmark specification (1.12) pass successfully the several robustness tests.

Besides, our results reveal that the highest impact of exchange rate was recorded in Italy, Spain, Greece and Portugal which is consistent with tendencies suggested by the statistical correlation reported in Section 3. It is worth stressing that these countries have historically a path of higher level of inflation. On the other hand, we find that the lowest ERPT coefficients are in Austria, Germany, Finland and France which are known as low-inflation regime countries.²⁰

To give further insights on the role of inflation regime, we can explore the expected positive link between the degree of ERPT and the inflation environment as argued by TAYLOR (2000). For illustrative purposes, we plot the ERPT elasticities against the mean of inflation rate for each country. Inflation is computed as the quarterly year-on-year changes of consumer prices index. In Figure 1.4, we report the correlation between pass-through and inflation average over 1990-2010. Initially, we have excluded Greece from the plot due to the relative high inflation level (7%) during this period. A simple visual inspection of Figure 1.4 reveal a clear positive relationship in line with Taylor's hypothesis. A weak degree of pass-through is associated with lower inflation rate. While countries with high-inflation environment, would experience higher degree of pass-through. This result is robust to the inclusion of Greece (see upper left subfigure in Figure A.1 in Appendix A.4). Furthermore, when considering the past inflation (inflation history) in EA countries, i.e. over 1990-1998 or 1979-1989, the positive correlation is still robust (see subfigures in Figure A.1 in Appendix A.4). In all, our results support the view that more stable macroeconomic conditions would entail a lower degree of ERPT into import prices.

²⁰We mention that according to Table 1.1 the simple correlation between import-price inflation and the rate of exchange depreciation in Austria was close to zero with a wrong (negative) sign. While when estimating the equation (1.12) for Austria, we find a pass-through elasticity equal to 0.29%. This confirms the importance of using an economically meaningful specification rather than purely statistical relationship between exchange rates and import prices.

Figure 1.4: Correlation ERPT and inflation between 1990-2010 (Greece excluded)

Note: y-axis: ERPT to import prices estimated from equation (1.12) over 1990-2010; x-axis: average of inflation over the same estimation period.

It is important to note that *CAMPA and GOLDBERG (2002, 2005)* has reported a limited role of macroeconomic variables, such as inflation environment, in explaining the extent pass-through in their sample of 23 OECD countries. As emphasized by the authors, ERPT is influenced more by the product composition of a country's exports than by macroeconomic factors. As a matter of fact, the hypothesis that the responsiveness of prices to exchange rate fluctuations depends positively on inflation seems to bear more on pass-through to consumer prices than on pass-through to import prices (see e.g. *CHOUDHRI, FARUQUEE, and HAKURA (2005)*, *CA'ZORZI, KAHN, and SÁNCHEZ (2007)* and *GAGNON and IHRIG (2004)*). Nevertheless, we believe that the pricing decision of a foreign firm, and therefore the choice between LCP and PCP strategy, depend on the macroeconomic conditions in the destination market. Countries with stable monetary policies are more likely to have their currencies chosen for transaction invoicing, and hence more likely to have low import-price pass-through.

Finally, we can explore another potential determinant of ERPT which is the degree of openness of a country. Intuitively, it is expected that the rate of pass-through is positively correlated with the openness of an economy. The larger presence of imports and exports in an economy, the larger the pass-through coefficient. The extent of trade openness can be measured as the ratio of exports and imports to domestic income or computed as the import penetration ratio, i.e. the participation of foreign firms in the domestic economy, measured by the share of imports in domestic consumption. However, few are studies who provide a strong evidence in this sense. For instance, in his VAR study, MCCARTHY (2007) find a little evidence that openness is positively correlated with ERPT to consumer prices, while no evidence of a statistically significant positive relationship with ERPT to import prices.²¹

In our EA sample, we aim to ascertain whether more open countries would experience a higher ERPT into import prices. The degree of trade openness is computed as the share of imports of goods and services in GDP.²² Besides, it is known that since the creation of the single currency, the share of trade affected by exchange rate fluctuations has been changed. Therefore, for more relevancy, on one hand, we plot the correlation of ERPT with (total) imports share over 1990-1998; on the other hand, the correlation is set out with respect to the extra-EA imports share over 1999-2010. In Figure 1.6, we report both total imports and imports coming from outside the EA as a share of GDP. It is important to note that there is a wide dispersion in terms of degree of openness in our sample. For the total imports share over 1990-1998, Belgium has the highest openness while Greece has the lowest. When considering the extra-EA imports over 1999-2010, the larger share is found in Netherlands, while the lowest is recorded in Portugal. We see that the inception of the euro has constituted a changing in the part of trade exposed to exchange rate fluctuations which may have a consequence on the ERPT behavior after the creation of the euro zone in 1999.

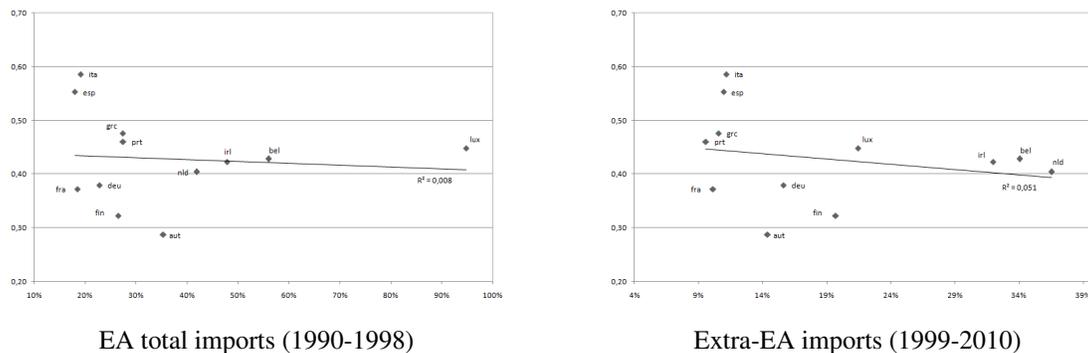
Tuning to the relationship between ERPT and openness as reported in Figure 1.5. A cursory look shows that the statistical correlation is close to zero with a slight

²¹CHOUDHRI, FARUQUEE, and HAKURA (2005), CA'ZORZI, KAHN, and SÁNCHEZ (2007) and GAGNON and IHRIG (2004) found no statistical link between pass-through to consumer prices and openness.

²²The data on the ratio of imports of goods and services to GDP are obtained from Eurostat and OECD's *Economic Outlook*.

negative sign. A higher import share as proxy for the degree of openness does not seem to be associated with a higher extent of ERPT. As mentioned above, the presence of a positive link between import openness and pass-through finds only weak empirical support. One potential explanation is that greater imports penetration may imply higher degree of competition for market share, thus implying lower ERPT. In fact, as mentioned by GUST, LEDUC, and VIGFUSSON (2010), the process of international globalization leading to high share of traded goods and high import content would induce a fall in pass-through. Following this reasoning, the authors explain that the higher trade integration has reduced the market power of U.S. producers at home and squeezed their U.S. profit margins.

Figure 1.5: Correlation between ERPT and degree of openness



Note: y-axis: ERPT to import prices estimated from equation (1.12) over 1990-2010; x-axis: ratio of imports to GDP.

Along with this vein, MARAZZI et al. (2005) explain that the increasing presence of China's exports in the U.S. market may also be partly responsible for the low levels of observed pass-through in the American economy in recent years. Especially, competition from Chinese firms may have constrained exporters from other countries from raising their prices in response to the dollar's depreciation, leading to lower degree ERPT than expected. Given these arguments, it is not surprising to find no evidence of strong association between pass-through into import prices and degree of trade openness.

7. Stability of ERPT Elasticities

In this section, we raise the question of whether the ERPT has changed over time in EA countries. Several macro studies have focused on the issue of the widespread and on-going decline in the pass-through. This decline has received more attention since it has important implications for the conduct and design of monetary policy. A frequently cited example include the case of some industrialized countries, namely Canada, Finland, Sweden and United Kingdom, which experienced a considerable depreciation of the exchange rate in the 1990s without consumer prices being affected as much as expected. This common experience has led to the widely held belief that pass-through of exchange rate changes into domestic inflation have declined in many these countries since the 1990s. For our country sample, there are many reasons to expect a changing in the ERPT behaviour. Especially, the formation of the EA would entail a change in macroeconomic environment and in the competitive conditions (by increasing the share of goods denominated in the single currency), and thus the extent of exchange rate transmission would be affected accordingly. Therefore, it is natural to ask whether the launch of the monetary union in 1999 constitute a break date in the pass-through mechanism across the EA countries.

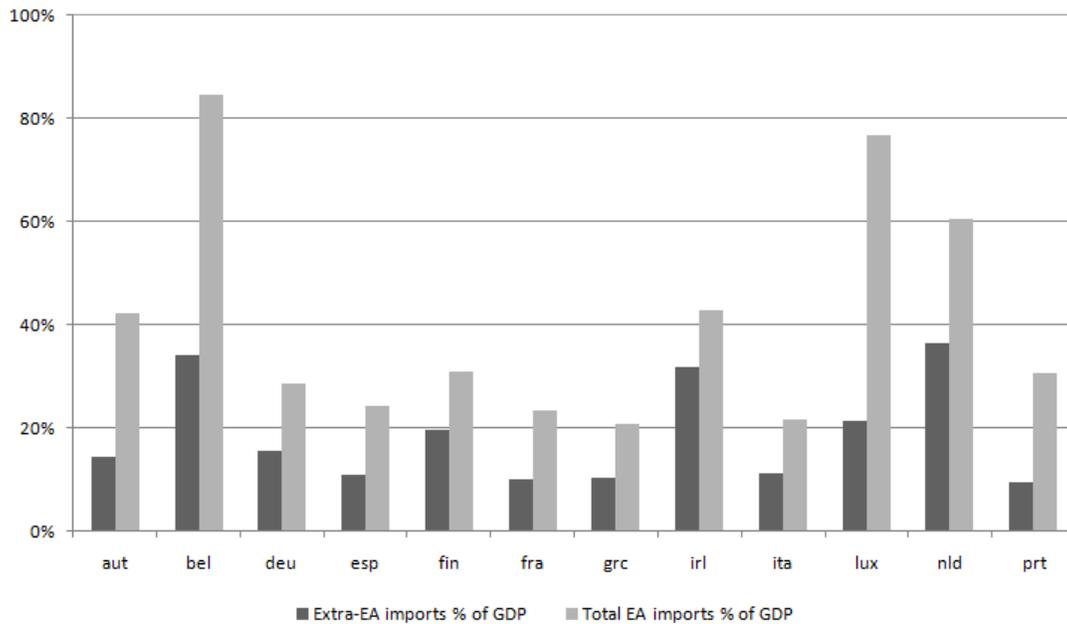
7.1. Is there a structural break around 1999?

A number of empirical studies has tested for the presence of structural break around the date of the inception of the euro. Using panel cointegration approach, DE BANDT, BANERJEE, and KOZLUK (2008) provide an evidence of a change around of the introduction of the common currency (1998-1999) or in the vicinity of the starting of the euro appreciation against the U.S. dollar (2001-2002). However, CAMPA and GOLDBERG (2005) and CAMPA and GONZÁLEZ (2006) provided a weak evidence in favour of the existence of a structural break around that time.

There are number of factors that may lead to a change in the rate of ERPT. As mentioned above, the proportion of trade exposed to exchange rate movements has diminished after the adoption of the single currency, and this has altered the magnitude of degree of openness in the respective EA countries. For example, as showed by

Figure 1.6, Portugal was more open to trade than Germany over 1990-1998, while since the starting of the monetary union, Portugal becomes less open than Germany. Such developments may lead to a change in the transmission of exchange rate movements. As explained by DORNBUSCH (1987), pass-through may be higher if the exporters are large in number relative to the presence of local competitors. However, as we have seen, the advent of the euro has reduced the market power of foreign firms relative to their domestic counterparts, and this would entail a decline in the responsiveness of import prices.

Figure 1.6: The share of imports in GDP (1999-2009)



Source: Eurostat and OECD's *Economic Outlook*.

Moreover, the choice of the currency of invoicing would be affected following 1999. It is expected that the share of trade being denominated in the euro would increase. As explained by DEVEREUX, ENGEL, and TILLE (2003), to the extent that the single currency becomes as the currency of denomination of trade for EA countries, ERPT elasticities would tend to reduce. To give a further insight on the expansion of the euro as an invoicing currency across some EA countries, in Table 1.3, we give the share of imports stemming from outside the EA with prices denominated in euro. We denote a

general increased use of the euro as the currency of denomination as it becomes a well established currency (mainly since 2002). For instance, MARAZZI et al. (2005) found that 1997 corresponds to the year after which the decline in U.S. import-price pass-through sped up. Given the large trade flows with Asian countries, the authors argued that the Asian financial crisis of 1997 have played a substantial role in the reduction of the pass-through to import prices. They also provide evidence suggesting that rising of exports from China to the American may also be partly responsible for the low levels of observed pass-through in recent years.

Table 1.3: The share of the euro as an invoicing currency of EA trade with the rest of the world (%)

Country	Imports of goods										
	2001	2002	2003	2004	2005	2006	2007	2008	2009	2010	2011
Belgium	47.2	53.7	57.8	55.7	51.2	58.3	56.1	56.4	57.7	53.0	55.7
Spain	49.7	55.9	61.1	61.3	56.0	54.8	56.7	58.8	60.6	59.1	57.7
France	42.6	40.8	44.1	45.7	46.3	44.7	44.8	44.2	44.3	45.5	53.0
Greece	29.3	35.4	39.2	39.6	32.6	32.3	33.6	37.3	37.9	30.8	33.1
Italy	40.8	44.2	44.5	41.2	39.4	43.0	44.3	47.8	49.7	46.9	-
Luxembourg	47.2	31.9	41.9	50.0	43.8	38.8	37.9	38.8	55.3	55.0	48.7
Portugal	50.3	54.9	58.1	58.0	54.4	52.6	51.8	53.7	56.6	52.1	45.7

Source: Review of the international role of the euro, European Central Bank, July 2012.

Therefore, to test for the possible decline in ERPT, as suggested by the above arguments, we perform tests of structural stability in the pass-through rates around the starting of the third stage of EMU, i.e. in the vicinity of 1999. To achieve this, we follow CAMPA and GOLDBERG (2005) and CAMPA and GONZÀLEZ (2006) by performing two types of structural change tests on the pass-through coefficients. First, we perform Chow tests assuming an exogenously imposed break point in 1999 or close to that date. In a second set of tests we allow for endogenously determined structural break points. It is possible that a change in ERPT elasticities does not happen on an exact date of 1999, thus, the Chow tests are also conducted for time break around the introduction of the euro.²³ Second, to check for the existence of an endogenous break any time over our sample period (1990:1-2010:4), we use ANDREWS (1993) and ANDREWS and

²³CAMPA and GOLDBERG (2005) and CAMPA and GONZÀLEZ (2006) assume that structural break might occur in May 1998, the month on which the parities among currencies replaced by the euro were announced.

PLOBERGER (1994) (AP hereafter) tests without specifying a priori the date at which the change in ERPT relationship takes place.

The results for the different of structural break tests are summarized in Table 1.4.²⁴ As for Chow tests, we are not able to reject the null of no structural break for 9 out of 12 EA countries. For these countries, the creation of the monetary union does not affect the extent of pass-through. Only for Belgium, Greece and Ireland the hypothesis of structural stability is rejected, implying that the formation of the euro area to have caused a change in the exchange rate transmission. Likewise, when applying ANDREWS (1993) and AP tests, there is a weak evidence in favour of the existence of a (statistically significant) structural break in ERPT into import prices across EA countries. We find endogenous breaks in the end of 1997 for Belgium and Italy and in the vicinity of 1998 for Greece and Ireland. We must be careful in our interpretation of these break points. As explained by CAMPA and GONZÁLEZ (2006), the change in ERPT elasticities around 1997-1998 is likely to be related to the negative oil price shock at that time rather than having any link to the formation of the euro zone. Generally speaking, we can say that the presence of a structural break in ERPT coefficients around 1999 does not occur systematically across EA countries.

Table 1.4: Structural break tests on ERPT elasticities

	Austria	Belgium	Germany	Spain	Finland	France
Chow test	0,201 (0,904)	10,183 (0,006)	0,190 (0,827)	0,587 (0,556)	1,819 (0,162)	0,062 (0,940)
ANDREWS (1993)	1,366 (0,938)	8,387 (0,055)	2,980 (0,558)	2,310 (0,710)	2,651 (0,630)	0,967 (0,971)
AP test	0,222 (0,758)	2,636 (0,024)	0,843 (0,249)	0,372 (0,559)	0,640 (0,346)	0,129 (0,949)
Break date	-	1997:04	-	-	-	-
	Greece	Ireland	Italy	Luxembourg	Netherlands	Portugal
Chow test	15,208 (0,000)	5,500 (0,064)	0,741 (0,690)	1,680 (0,432)	0,836 (0,658)	2,459 (0,293)
ANDREWS (1993)	8,601 (0,050)	3,898 (0,390)	6,668 (0,120)	4,536 (0,301)	1,974 (0,792)	3,077 (0,538)
AP test	2,818 (0,019)	1,177 (0,153)	1,488 (0,100)	0,736 (0,295)	0,280 (0,671)	0,609 (0,365)
Break date	1998:02	1998:03	2007:04	-	-	-

Note: Numbers in in parenthesis p -value of the tests. As test statistic, ANDREWS (1993) uses the maximum of the LM statistics, while AP test (ANDREWS and PLOBERGER (1994)) uses the geometric mean.

²⁴Greece joined the monetary union in 2001, so Chow test is preformed around this date.

It is noteworthy that a change in ERPT may not happen exactly in a particular point in time, such as 1999. The decline in exchange rate transmission may be gradual rather than associated with a distinct break date. Furthermore, as discussed by DE BANDT, BANERJEE, and KOZLUK (2008), the changing behaviour in pass-through mechanism may have started before the date of the creation of the euro (for example during the first or the second stage of EMU) or after the strengthening of the common currency since 2002. For instance, the acceptance of the euro as an invoicing currency may be gradual and thus picked up with a lag as the euro became a well established. DE BANDT, BANERJEE, and KOZLUK (2008) found that the appreciation of the euro against U.S. dollar since in 2002, has caused a change in long-run relationship of ERPT.

Otherwise, as we have seen before, there has been a dramatic fall in inflation levels during 1990s compared to 1980s in most of the EA countries (see Figure 1.1). In fact, the process of EMU has entailed some convergence of average inflation rates across the EA members, as a result of efforts to fulfill Maastricht convergence criteria. Thus, the reduction in inflation rates has started largely before the inception of the euro. Given that inflation environment is an important macro determinant of ERPT, one can think that the shift towards more credible and anti-inflationary monetary policy regimes may contribute to lowering the response of import prices to currency movements in EA. Drawing on this intuition, it is expected that the extent of pass-through was higher in the 1980s than during the last two decades (1990-2010). This would be especially the case of EA countries with historically higher inflation levels, namely Greece, Ireland, Italy, Portugal and Spain. In the next sub-section, we try to estimate the ERPT over the 1980s and to compare results with those obtained over 1990-2010.

7.2. ERPT in the 1980s

A recurrent exercise in the empirical literature is to estimate the ERPT over different sub-sample periods, to test for the conventional wisdom of the decline of pass-through.²⁵ Given the steady decline in inflation rates in our sample of EA, we aim to investigate

²⁵For instance, the split-sample approach was used by GAGNON and IHRIG (2004) for 20 industrialized countries between 1971 and 2000. The authors have estimated the transmission of exchange rate over two sub-samples periods, with break dates chosen based on the observed behaviour of inflation. Thus, the first subsample period is a period of high inflation environment, while the second subsample has lower and

whether this changing in macroeconomic environment has fostered the decline in the ERPT. Therefore, we reestimate our benchmark model (1.12) over 1979:2-1990:2, i.e. before the inception of the first stage of EMU, and compare the pass-through elasticities with those obtained over 1990:3-2010:4.

As reported in Table 1.5, we point out more pronounced cross-differences in ERPT than recorded over 1990:3-2010:4. There was divergent macroeconomic conditions across EA countries during the 1980s, especially between peripheral and core economies. Thereby, it is expected that the general process of European convergence, which has started before the introduction of the Euro in 1999, would entail a reduction in the variability of pass-through within EA members states. Also, we note that the hypothesis of null ERPT was rejected for all countries in our sample, while the full ERPT hypothesis is accepted only for Spain. For the latter country, we denote a higher responsiveness of import prices, i.e. when the rate of depreciation increase by 1%, the Spain import-price inflation rises by 0.95%. The smallest rate of pass-through is found in Luxembourg, where a one percent rising in exchange rate depreciation lead to increase in the rate of inflation of import prices by 0.19 percent.

When comparing elasticities estimated in the 1980s (over 1979:2-1990:2) with those in the last two decades (over 1990:3-2010:4), we found a general decline in the rate of pass-through in most of EA countries, except for Belgium and Luxembourg (see Figure 1.7). On average, the import-price pass-through fell from 0.54% over the 1980s to 0.43% over 1990-2010, which corresponds to a decrease of about 0.14% on average.²⁶ In their sample of 23 OECD countries, CAMPA and GOLDBERG (2005) compared ERPT estimated over 1975-1989 with those over 1990-2003 and found that short- and long-run ERPT elasticities declined for 15 out of 21 countries and increased for the other 6 countries. On average, the decline in the short-run import price pass-through is about

more stable inflation. The authors found a strong decline in the pass-through across the two time periods and conclude that is due to an increased emphasis of monetary policy on stabilizing inflation.

²⁶To provide a statistical significance of our results, we display the point estimates of ERPT with 95% confidence intervals over the two periods in Figure A.2 in Appendix A.5. We see that the decline is more pronounced, especially for Spain, Finland and France, with the rates of pass-through are strongly different between the two sample periods. It is interesting to note that Spain has had a prior history of high inflation, namely double-digit inflation rate during the 1970s and the 1980s, while in the last two decade the increase in CPI has not exceed the 5% on average. We can expect that this shift towards a stable inflation regime has contributed to the lowering of the Spanish pass-through.

0.04 in CAMPA and GOLDBERG (2005), however, the average fall reported in our study is three times as large.

Table 1.5: Estimation results over 1979:2-1990:2

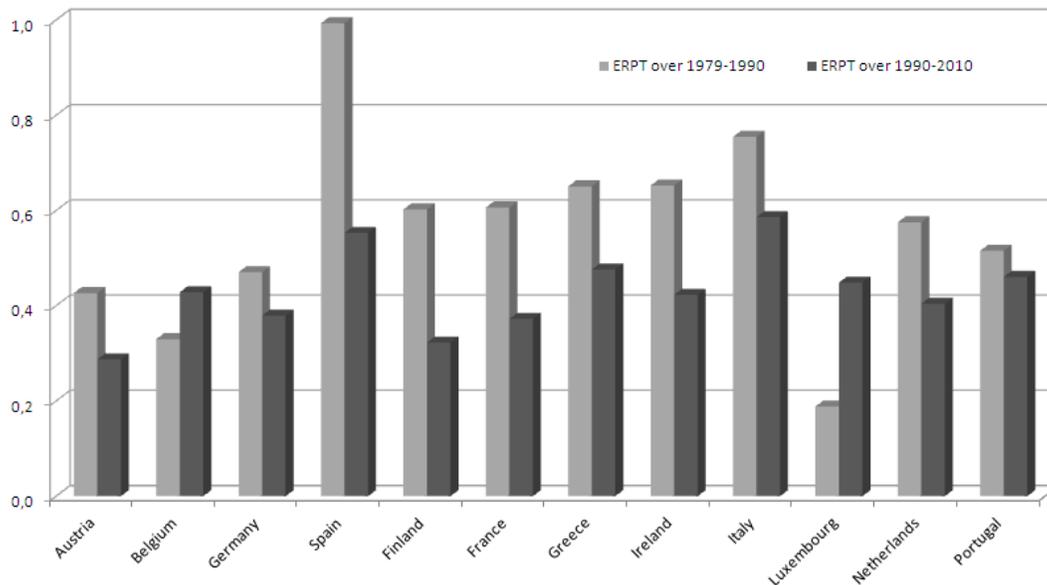
	Austria	Belgium	Germany	Spain	Finland	France
Constant	0,007 (0,118)	0,009 (0,010)	0,002 (0,622)	0,027 (0,003)	0,008 (0,108)	0,006 (0,145)
Δe_t	0,427 (0,000) [0,000]	0,330 (0,000) [0,000]	0,470 (0,000) [0,000]	0,993 (0,000) [0,950]	0,602 (0,000) [0,000]	0,606 (0,000) [0,000]
Δw_t^*	0,615 (0,000)	0,309 (0,000)	0,661 (0,000)	1,314 (0,000)	0,765 (0,000)	0,773 (0,000)
gap_t	0,099 (0,699)	0,570 (0,000)	-0,113 (0,509)	0,684 (0,134)	0,040 (0,714)	0,528 (0,007)
Observations	45	45	45	45	45	45
R^2	0,712	0,636	0,747	0,747	0,594	0,772
Wald test	3,567	0,030	6,178	3,332	5,504	2,935
p -value	(0,067)	(0,864)	(0,017)	(0,076)	(0,024)	(0,095)
Chow test	3,419	3,631	18.23721	2,732	5,475	8,510
p -value	(0,033)	(0,026)	(0,000)	(0,065)	(0,004)	(0,000)
	Greece	Ireland	Italy	Luxembourg	Netherlands	Portugal
Constant	0,021 (0,006)	-0,003 (0,574)	0,003 (0,685)	0,010 (0,001)	0,002 (0,791)	0,024 (0,001)
Δe_t	0,650 (0,000) [0,000]	0,652 (0,000) [0,000]	0,755 (0,000) [0,015]	0,188 (0,000) [0,000]	0,575 (0,000) [0,000]	0,515 (0,000) [0,000]
Δw_t^*	0,821 (0,000)	0,884 (0,000)	0,897 (0,000)	0,160 (0,064)	0,774 (0,000)	0,631 (0,002)
gap_t	-0,051 (0,712)	0,109 (0,571)	0,225 (0,492)	-0,067 (0,518)	-0,107 (0,684)	0,271 (0,112)
Observations	45	45	45	45	45	45
R^2	0,679	0,685	0,692	0,494	0,631	0,520
Wald test	1,569	3,768	0,729	0,189	1,649	0,643
p -value	(0,218)	(0,060)	(0,398)	(0,667)	(0,207)	(0,428)
Chow test	8,551	2,087	2,924	0,918	2,387	13,899
p -value	(0,000)	(0,124)	(0,054)	(0,400)	(0,092)	(0,000)

Note: Estimation are based on equation (1.12) over 1979:1-1990:2. Numbers in parentheses are p -values. For the exchange rate coefficient, p -values in parentheses are based on the null hypothesis of zero ERPT, i.e. $H_0: \beta_1 = 0$, while p -values in square brackets corresponds to the null of full ERPT, i.e. $H_0: \beta_1 = 1$. Wald test is performed for $H_0: \beta_1 - \beta_2 = 0$. Chow test is performed for the hypothesis that a structural break took place around 1990.

Moreover, it is interesting to note that when we perform Chow tests assuming an exogenously imposed break point around 1990, we found the null of ERPT stability are strongly rejected for most of EA countries, except for Ireland and Luxembourg (see last rows in Table 1.5. These results appear overall supportive of the hypothesis of a change

in ERPT mechanism over time. Although the change is not statistically significant for some EA countries, as reported in Figure A.2, we can say that there has been a tendency toward declines in pass-through in our sample. The fact that the behavior of pass-through in the last two decades has been different than was the case before seems compelling.

Figure 1.7: Decline of ERPT into import prices



Source: Personal calculation.

7.3. Evidence from rolling regressions

To give further evidence on the significant decline in ERPT in our country sample, in this section, we use a rolling window regression approach. This allows us to check how pass-through has changed over the time. For this purpose, ERPT elasticities will be estimated from equation (1.12) with a 10-year moving window rolled forward one quarter at a time. We start with the window 1979:2- 1989:1 and finish with 2001:1-2010:4. This will trace the evolution of the responsiveness of import prices in EA countries.²⁷

²⁷As robustness tests, we followed IHRIG, MARAZZI, and ROTHENBERG (2006) by considering different sample windows, of 15 years for example, in addition to a 10-year window. These robustness tests are important, because without them it is not clear whether a change in the pass-through coefficient reflects the new quarters of data entering the sample or the old quarters of data dropping out of the sample.

The rolling estimates of import-price pass-through are shown in Figure 1.8 (estimates with standard error bands are reported in Figure A.3 in Appendix A.6). Also, we have reported inflation rates on the same plots to assess whether the shift towards stable inflation environment was synchronous to the decline in ERPT. For a better understanding of plots in Figure 1.8, the first observation which lies above 1989:1 (on the horizontal axis) corresponds to the first 10-year sample, i.e. the period 1979:2-1989:1.²⁸ While the latest 10-year sample, i.e. the period 2001:1-2010:4, is reported as 2010:4 on the horizontal axis.

A careful inspection of Figure 1.8 reveals that ERPT to import prices was higher during the 1980s (in the first 10-year window) but appears to trend down afterward in most of EA countries, except for Belgium and Luxembourg. The degree of pass-through decreased between our earliest and latest 10-year samples.²⁹ For example, in France, the exchange-rate sensitivity of import prices was more than 0.60% in the 1980s, while it has begun a steady decline since 1994 to reach 30% of pass-through by the end of 1996. It is interesting to note that pass-through has been high until the end of the Exchange Rate Mechanism (ERM) crisis of the European Monetary System in the beginning of the nineties (1992-1993), a period during which many European currencies have experienced substantial depreciations³⁰. Since the launch of second stage of EMU in 1994, there is a strong evidence of lowering ERPT for the most of the EA members. This decline came after the implementation of the Maastricht treaty which emphasize on the achievement of a high degree of price stability (among other convergence criteria).³¹

In doing this, we find that the size of the sample window does not really matter. Our results are still robust since the decline in ERPT is apparent in most of our country sample.

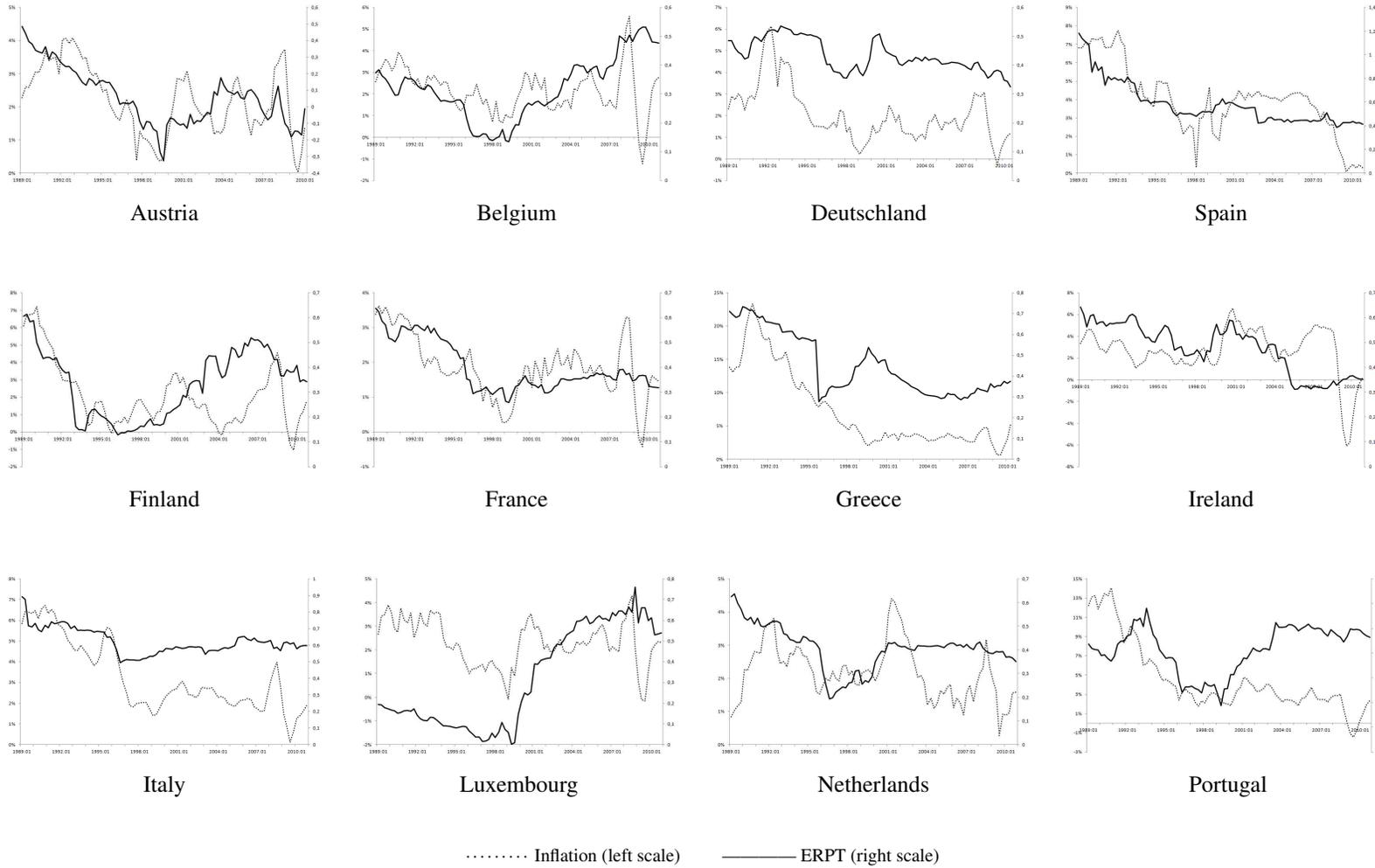
²⁸The estimates obtained from the first 10-year sample should be close to those displayed in Table 1.5 in the previous sub-section.

²⁹For some countries, the decline is not significant at 95 percent confidence interval.

³⁰For example, Italy left the ERM in September 1992.

³¹As stipulated the Maastricht convergence criteria, each country's inflation in 1997 had to be less than 1.5 percentage points above the average rate of the three European countries with the lowest inflation over the previous year.

Figure 1.8: Moving Window ERPT and Inflation in the EA



Our results are in line with IHRIG, MARAZZI, and ROTHENBERG (2006) who estimates import-price pass-through in G-7 countries using a rolling regression framework. For France, they reported that ERPT was about 0.50% and stable through 1996, while in 1997, the estimate has started to fall to reach less than 0.2% in the end of 2004. Among the G-7 countries, the authors found that has the lowest level of import-price pass-through in the end of sample. However, IHRIG, MARAZZI, and ROTHENBERG (2006) explained that this lowering in the rate of pass-through might be correlated the 1997 Asian financial crisis. As discussed in MARAZZI et al. (2005), this explanation would be appropriate for the U.S case. Knowing that a substantial portion of U.S. imports come from Asia, it is expected that the Asian crisis of 1997 may have contributed importantly to the decline in pass-through in the U.S. The authors also provide evidence suggesting that the rising prominence of competition from China may also be partly responsible for the low levels of U.S. ERPT. For Germany, they reported a slight decline in the estimates as in our study. With an already low level of pass-through, it is expected that German import-price pass-through did not fall very much. As noted before, inflation levels have fallen markedly in most of EA countries since the beginning the 1990s (see Figure 1.1). It is worth highlighting that this shift towards stable low-inflation regime has coincided with a general decline in the extent of pass-through.

The visual inspection of Figure 1.8, show that there is a broad downward tendency for both inflation and ERPT. This would corroborate the positive correlation between inflation and ERPT as reported in Section 6. Finally, we can note that the wide swings of the single currency during the first three years of the monetary union was a serious threat to price stability in the EA by putting upward pressure on import costs and producer prices.³² Plots in Figure 1.8 confirm the rising in inflation rates in most of EA Members States from 1999 to 2000 due to the extensive depreciation of the euro. It should be noted that this outcome would explain why ERPT has risen in Belgium and Luxembourg instead of decreasing. For the latter countries, the inflation levels were already low, and it is not surprising that pass-through would increase in accordance with the rising of inflationary pressures at the beginning of the EA. Overall, we can say that exchange rate changes continue to lead to a significant pressures on domestic prices, justifying the growing interest in the issue of pass-through in the context of the EA.

³²See ECB statements by F. Duisenberg (President of the European Central Bank) in 2000.

8. Factors influencing ERPT: Evidence from dynamic panel data

So far, our ERPT estimates are based on a static model which corresponds to equation (1.12). However, it is known that the responses of import prices to exchange rate changes may not be fully manifested instantaneously, especially when foreign firms take time to adjust their prices in the domestic currency. Thus, as emphasized by some empirical studies, it is important to account for the potential inertial behaviour of import prices by estimating a dynamic model (see e.g. BUSSIÈRE, 2012; OLIVEI, 2002; YANG, 2007, among others).³³ This is typically accomplished by including lagged import prices as an explanatory variable. This allows for the possibility of delayed adjustment of domestic-currency import prices. Furthermore, in this section, we aim to estimate an aggregate ERPT into import prices for the whole EA. To achieve this, we call for the use of a panel data framework.

Besides, in the previous section, we have tested for the influence of some factors or events on the extent of pass-through. We have found that inflation environment play a substantial role while other factors, such as the creation of the EA and the degree of openness, has no significant effect on the ERPT. In this section, we try to provide further insights on the factors influencing the transmission of the exchange rate by employing a different econometric approach, i.e. dynamic panel data model. Other important factors are related to the cycles and large exchange rate movements that occurred many times in recent years across EA countries (see Figure 1.9). As we have seen in the previous section, the large swings of the euro since 1999 might affect the price stability in the monetary union. Also, we have noted that pass-through has been higher until the end of the ERM crisis, an episode during which many European currencies have experienced substantial depreciations. Following this instability period, we have provided a strong evidence of a decline in the ERPT for the most of the EA members. To check whether these events have impacted the responsiveness of import prices, we construct two dummy variables and include them in our dynamic panel data model interacted with the exchange rate:

³³Other models (e.g. MARAZZI et al., 2005) do not include the lagged dependent variable but include more lags of the explanatory variables instead.

- D^{ERM} is a dummy variable that takes the value of one during the time of ERM crisis (1992-1993), and zero otherwise.³⁴
- D^{Dep} is a dummy variable taking on the value one during the first three years of the creation of the euro (1999-2001), and zero otherwise.

To ensure the robustness of our previous findings with respect to the creation of the euro and the inflation environment, as potential factors influencing the pass-through, the following dummy variables will be created:

- D^{Euro} is a dummy variable that takes the value of one since 1999 (the date of the inception of the euro), and zero before.
- $D^{Inflation}$ is a dummy variable that takes the value of one during low inflation periods, and zero otherwise.

To identify low-inflation periods, we follow the approach of BAILLIU and FUJII (2004) by implementing BAI and PERRON (2003) multiple break test on quarterly inflation series in each country. Once one break or more are identified, together with visual inspection of inflation series, we can define the periods of high and low inflation, respectively.³⁵ The results of BAI and PERRON (2003) multiple break test are summarized in Table A.6 in Appendix A.7. Also, plots depicting the inflation series for each EA country are reported in Figure A.4 in Appendix A.7 with vertical lines representing the dates at which the structural breaks were identified. According to these results, we found evidence of at least one break in all EA countries, and for some countries, two breaks were identified. It is interesting to note that for several EA countries periods of low inflation lie between the end of ERM crisis (in 1994) and the launch of the euro (in 1999). It stands out that during the first years of the monetary union there was an uprising of inflation levels across our country sample.

³⁴Since they joined the ERM mechanism in a later date, Austria, Finland and Greece are excluded from the estimation when we consider this dummy variable.

³⁵To test whether pass-through declined following a change in the inflation regime, GAGNON and IHRIG (2004) split their sample between high and low inflation period without testing for the presence of structural breaks in inflation series. However, our study follows BAILLIU and FUJII (2004) by formally testing for structural breaks to identify low-inflation periods in our EA countries.

Therefore, to gauge the importance of these different factors, we modify our benchmark pass-through equation (1.12) to have all the elements of a dynamic panel data model as follows:

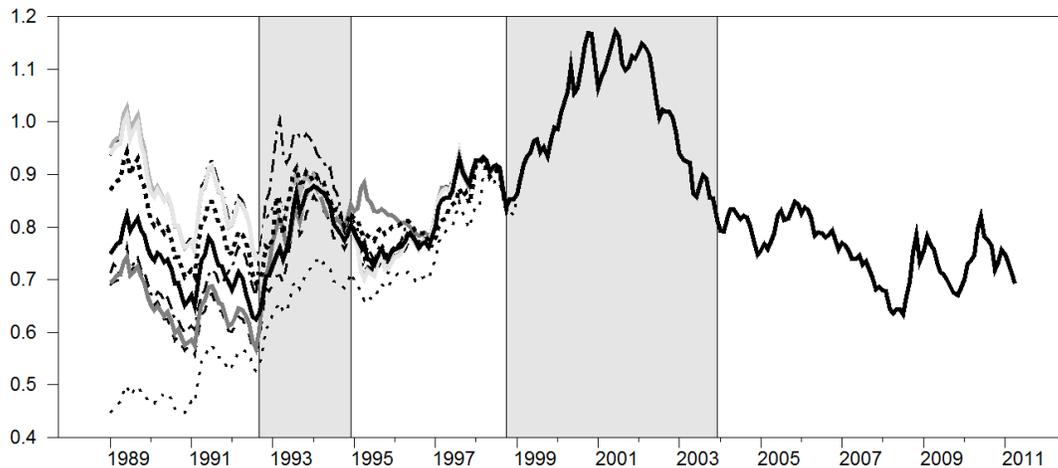
$$\Delta p_{i,t}^m = \alpha_i + \beta_1 \Delta e_{i,t} + \beta_2 \Delta w_{i,t}^* + \beta_3 gap_{i,t} + \beta_4 \Delta p_{i,t-1}^m + \beta_5 D^j \times \Delta e_{i,t} + \varepsilon_{i,t} \quad (1.13)$$

where α_i is a country-specific effect and D^j is the dummy variable (interacted with the exchange rate depreciation $\Delta e_{i,t}$) chosen in the set $\{D^{ERM}, D^{Dep}, D^{Euro}, D^{Inflation}\}$ to captures specific events, such as the ERM crisis (D^{ERM}) and the large depreciation of 1999-2001 (D^{Dep}), or shifts in macroeconomic environment, such as the introduction of the euro (D^{Euro}) and the low-inflation regime ($D^{Inflation}$). Our panel data set consists of annual observations for our 12 EA countries over 1990-2010. We estimate a separate regression for each dummy variable D^j , where $j = ERM, Dep, Euro, Inflation$. In our dynamic specification (1.13), it is possible to estimate the immediate effect of the exchange rate on import prices, i.e. the short-run ERPT, given by the coefficient β_1 . Also, due to the lagged adjustment of import-price inflation, we can compute the long-run ERPT given by $\beta_1/(1 - \beta_4)$. To check whether our interactive terms ($D^j \times \Delta e_{i,t}$) has an impact on the extent of pass-through, we compute the short-run ERPT as $(\beta_1 + \beta_5)$ and long-run ERPT as $(\beta_1 + \beta_5)/(1 - \beta_4)$, respectively.

Since, we opt for a dynamic specification, we use a generalized method of moments (GMM) estimator for dynamic panel data models to estimate the aggregate ERPT in the EA. The dynamic structure of our benchmark specification causes OLS estimators to be biased and inconsistent, since the lagged import prices is correlated with error term. ARELLANO and BOND (1991) propose a GMM procedure that is applied to the equation in first differences using a set of appropriate instruments to correct for the bias created by the presence of the lagged dependent variable as a regressor. This procedure is also suitable in a situation where one or more of the explanatory variables suspected to be endogenous rather than exogenous. In fact, exchange rate may be considered an endogenous variable as predicted by the purchasing power parity (PPP), and then causality may run in both directions (from exchange rate to prices and vice versa). Thus, using ARELLANO and BOND (1991) procedure, this endogeneity could be

treated by instrumenting for the exchange rate. When estimating the ERPT for the EA, we will explore both cases, i.e. whether the exchange rate is exogenous or endogenous.

Figure 1.9: European Currencies during the last two decades



Source: OCDE

An important drawback is that the GMM estimator is designed for short time dimension (T) and a larger country dimension (N). However, our panel data set panel has $T = 21$ and $N = 12$. To check the reliability of our GMM estimates, we provide also estimation results for the pooled OLS and fixed effects estimators. We note that the country-specific effect would be better modelled as fixed rather than random for two reasons: first, contrary to the random effects model, the estimation of a panel data model with fixed effects does not rely on the assumption that the unobservable individual effects must be uncorrelated with the explanatory variables; second, random effects estimation is relevant when the observations are drawn randomly from a given population. However, our sample contains a particular group of countries, which are the 12 EA Member States, and not a random sample from a larger group of countries. We mention that when using Arellano and Bond's dynamic panel-data GMM estimator, the country-specific effect is removed by the first-difference transformation.

Turning to our estimations results, we begin by estimating the equation (1.13) without including the interactive dummy variables to provide some insight on the

aggregate ERPT in the EA. Thus, the model to estimate has the following form:

$$\Delta p_{i,t}^m = \alpha_i + \beta_1 \Delta e_{i,t} + \beta_2 \Delta w_{i,t}^* + \beta_3 gap_{i,t} + \beta_4 \Delta p_{i,t-1}^m + \varepsilon_{i,t} \quad (1.14)$$

Results of different estimations techniques are displayed in Table 1.6. As explained before, we use two versions of the GMM estimator: GMM1 consider exchange rate as exogenous by instrumenting only for the lagged import prices; while GMM2 instruments for both the lagged dependent variable and the exchange rate. The methodology developed by ARELLANO and BOND (1991) assumes that there is no second order autocorrelation in the first-differenced errors. To validate this assumption, we test the null hypothesis of no autocorrelation using the *m2* test for autocorrelation. Additionally, to ensure the validity of the instruments, in the sense that they are not correlated with the errors in the first-differenced equation, the Sargan/Hansen test of over-identifying restrictions is performed. As reported in Table 1.6, the results of these two specification tests support the validity of the instruments for the GMM estimations.

As regards the estimation results, the coefficients of the key variables are statistically significant with expected signs in our panel of 12 EA countries. In contrast to the individual regressions in Section 5, we found output gap statistically significant. the output gap which is found to be positively significant only for 4 out of 12 EA countries. Very few EA countries have had a positive significant effect of the output gap on the import prices. Now, we see that our panel data framework has enhanced the reliability of the pass-through equation estimates. Furthermore, estimation results using the pooled OLS and fixed-effect estimators are by and large similar to the GMM estimations. As shown in Table 1.6, estimates are fairly robust across the different estimation techniques. Concerning pass-through estimates, the quarterly contemporaneous effect of the exchange rate movement, i.e. the short-run ERPT, is about 0.64% according to GMM1. The transmission is relatively high compared to the average of individual pass-through elasticity reported in Section 5 - which is equal to 0.43% - but incomplete, since the null of full ERPT is rejected throughout different estimation techniques (see *p*-values in square brackets for $H_0 : \beta_1 = 1$). It is possible that the frequency dimension of panel date set (annual instead of quarterly) and the use of dynamic model instead of static

one explains this difference in pass-through estimates. In the long-run, a 1% change in the rate of depreciation leads to 0.75% increase in the import-prices inflation in our EA sample. The long-run ERPT is slightly higher than in the short-run but still incomplete.

Table 1.6: Panel ERPT Estimates over 1990-2010

	GMM1	GMM2	Fixed Effects	Pooled OLS
Δpim_{t-1}	0,159 (0,155)	0,112 (0,000)	0,051 (0,219)	0,150 (0,002)
Δe_t	0,634 (0,000) [0,000]	0,646 (0,000) [0,000]	0,614 (0,000) [0,000]	0,623 (0,000) [0,000]
Δw_t^*	0,911 (0,000)	0,880 (0,000)	0,837 (0,000)	0,856 (0,000)
gap_t	0,249 (0,000)	0,209 (0,000)	0,245 (0,000)	0,240 (0,001)
Long-run ERPT	0,754 (0,000) [0,022]	0,728 (0,000) [0,000]	0,650 (0,000) [0,000]	0,734 (0,000) [0,000]
Observations	211	211	221	221
Sargan/Hansen test	10,840 (0,287)	9,500 (0,798)		
m_2 test for autocorrelation	0,370 (0,713)	0,420 (0,672)		
R^2			0,751	0,692

Note: Estimations are based on equation (1.14). Short-run ERPT corresponds to β_1 and long-run ERPT refers to $\beta_1/(1 - \beta_1)$. p -values in parentheses are based on the null hypothesis of zero ERPT, while p -values in square brackets corresponds to the null of full ERPT. The m_2 test for autocorrelation has a null hypothesis of no autocorrelation, while Sargan/Hansen test has the null hypothesis that model and over-identifying conditions are correct specified.

The estimated pass-through elasticities reported here are close to CAMPA and GONZÁLEZ (2006) with elasticities average 0.62 and 0.78 in the short- and long-run, respectively. However, in the long-run, the authors found that the hypothesis of complete pass-through was accepted for Spain, Finland, France, Greece, Italy and Portugal. Using dynamic panel data model for 11 industrialized countries, BAILLIU and FUJII (2004) suggested an exchange-rate sensitivity of import prices close to 0.75% in the short-run and near complete (0.91%) over the longer run. Overall, our results corroborate the conventional wisdom that the degree of ERPT is incomplete in the short-run. However, in the long-run, we found no evidence of complete pass-through in contrast to the existing literature. For our 12 EA countries, we conclude that partial ERPT is the best description for import price responsiveness in both short- and long-run.

As a next step in this section, we investigate whether some macroeconomics factors may influence the extent of pass-through. As discussed, these factors are the ERM crisis (D^{ERM}), the large depreciation of 1999-2001 (D^{Dep}), the introduction of the euro (D^{Euro}) and the low-inflation regime ($D^{Inflation}$). Thus, we estimate the equation 1.14 by including our interactive dummy variables of interest ($D^j \times \Delta e_{i,t}$) separately (as in equation(1.13)). The ERPT elasticity in each macroeconomic regime is equal to $(\beta_1 + \beta_5)$ in the short-run, while it corresponds to $(\beta_1 + \beta_5)/(1 - \beta_4)$ in the long-run. The use of interactive dummy variables to capture some the effect of some specific events was a typical approach in the empirical literature. For instance, BAILLIU and FUJII (2004) has constructed two policy dummy variables indicating shifts in the inflation environment in the 1980s and 1990s, to check the impact of shifting towards low-inflation regime on ERPT. Their results indicate that the decline in pass-through over time was brought about by the inflation stabilization episodes that took place in the 1990s rather than in the 1980s. In a similar vein, to test whether the adoption of inflation targeting has had an impact on the degree of ERPT to consumer prices, EDWARDS (2006) created a dummy variable that takes the value of one at the time of the adoption of the inflation targeting, and zero otherwise. Using quarterly data for the period 1985-2005 for seven countries - two advanced and five emerging - that have adopted inflation targeting, the author found that pass-through has declined in most of cases since the adoption of inflation target regime.

The results of the impact of our macroeconomic factors are summarized in Table 1.7. First, the ERM crisis over 1992-1993 seems to do not affect the sensitivity of import prices to exchange rate movements in the EA. The interactive term ($D^{ERM} \times \Delta e_{i,t}$) has a positive sign, meaning that ERPT is rising during this episode, but is not statistically significant different from zero throughout the different estimation techniques. Thus, when computing the short- and long-run pass-through, as displayed in Table 1.8, we found that estimates increase slightly in 1992-1993. Second, regarding the effect of the euro depreciation over 1999-2001, we point out a significant increasing in the responsiveness of import prices over this period. The interactive dummy variable ($D^{Dep} \times \Delta e_{i,t}$) is positively significant according to GMM estimators (GMM1 and GMM2). According to GMM1 estimations in Table 1.8, the ERPT coefficient increased

significantly from 0.62% to reach 0.70% in the short-run, while it changes from 0.70% to 0.78% in the long-run.

Table 1.7: Effects of macroeconomic environment on ERPT

Variables	GMM1	GMM2	Random Effects	Pooled OLS
Effect of the ERM crisis				
Δe_t	0,621 (0,000)	0,601 (0,000)	0,606 (0,000)	0,612 (0,000)
$\Delta e_t \times D^{ERM}$	0,065 (0,215)	0,069 (0,289)	0,023 (0,656)	0,031 (0,611)
Effect of euro depreciation (1999-2001)				
Δe_t	0,619 (0,000)	0,559 (0,000)	0,593 (0,000)	0,602 (0,000)
$\Delta e_t \times D^{Dep}$	0,075 (0,046)	0,118 (0,065)	0,074 (0,147)	0,073 (0,174)
Effect of the monetary union				
Δe_t	0,618 (0,000)	0,649 (0,000)	0,600 (0,000)	0,620 (0,000)
$\Delta e_t \times D^{Euro}$	-0,027 (0,167)	-0,037 (0,211)	0,031 (0,434)	0,037 (0,359)
Effect of low-inflation regime				
Δe_t	0,677 (0,000)	0,665 (0,000)	0,650 (0,000)	0,658 (0,000)
$\Delta e_t \times D^{Inflation}$	-0,078 (0,028)	-0,101 (0,024)	-0,084 (0,023)	-0,082 (0,061)

Note: Estimations are based on equation (1.13). Coefficients reported here are β_1 for Δe_t and β_5 for the interactive terms ($D^j \times \Delta e_t$). Numbers in parentheses are p -values.

This outcome is very important and has several implications. On one hand, our finding justify why the pass-through of exchange rate was a cause for concern for the ECB since the launch of the monetary union in 1999. The dramatic depreciation of the single currency was a serious threat to price stability over the first three years of the euro. On the other hand, this would explain the failure to find a decline in ERPT since the formation of the EA in the 1999 (see Section 7). Despite the reduction of share of imports affected by exchange rate fluctuations and the increase of the euro as an invoicing currency, we did not find that ERPT has declined since the adoption of the single currency. We think that the significant increase of the extent of pass-through in the beginning of the creation of the euro has prevented the expected decline of the sensitivity of the import prices to exchange rate changes since 1999. After several years of depreciation, the euro has started off on a relatively stable appreciation since

2002. Thus, to the extent that the euro became a well established currency, foreign firms would tend to choose it as the currency of denomination of their exports (LCP strategy), leading to a lesser degree of pass-through. This outcome was confirmed by DE BANDT, BANERJEE, and KOZLUK (2008) who reported significant changing in the ERPT behaviour in the vicinity of the strengthening of the euro since 2002.

As consequence, it is not surprising that the interactive terms included to capture the effect of introduction of the euro are found to be insignificant. As showed in Table 1.7, $(D^{Euro} \times \Delta e_{i,t})$ has a non-significant negative effect by the GMM estimators, while a non-significant positive coefficients are found with fixed effects and pooled OLS estimators. These findings corroborate what we find in the previous section, that is, the inception of the monetary union in 1999 does not entail a changing in the magnitude of the ERPT in the EA.

Table 1.8: Short- and long-run ERPT over different macro environments

	Short-run ERPT	Long-run ERPT
Outside the EMS crisis	0,621 (0,000) [0,000]	0,709 (0,000) [0,012]
During the EMS crisis	0,686 (0,000) [0,000]	0,783 (0,000) [0,002]
Outside the depreciation 1999-2003	0,619 (0,000) [0,000]	0,694 (0,000) [0,000]
During the depreciation 1999-2003	0,694 (0,000) [0,000]	0,778 (0,000) [0,000]
Pre-EA	0,618 (0,000) [0,000]	0,883 (0,000) [0,153]
Post-EA	0,591 (0,000) [0,000]	0,844 (0,000) [0,023]
During high inflation	0,677 (0,000) [0,000]	0,753 (0,000) [0,015]
During low inflation	0,598 (0,000) [0,000]	0,666 (0,000) [0,012]

Note: Estimations are based on equation (1.13) using GMM1 method. The ERPT elasticity in each macroeconomic environment is equal to $(\beta_1 + \beta_5)$ in the short-run, while it corresponds to $(\beta_1 + \beta_5)/(1 - \beta_4)$ in the long-run. p -values in parentheses are based on the null hypothesis of zero ERPT, while p -values in square brackets corresponds to the null of full ERPT.

Finally, we investigate whether the inflation environment indeed influence the extent to which the exchange rate changes are transmitted into import prices. It stands out from 1.7 that the low-inflation periods, as given by BAI and PERRON (2003) multiple break test, dampen significantly the effect of exchange rate changes. The estimated coefficient for $(D^{Euro} \times \Delta e_{i,t})$ is negative and statistically significant throughout the different estimation techniques. The short-run pass-through rate is roughly 0.70% during high-inflation periods, and is reduced to around 0.60% following a change towards more stable inflation environment. According to long-run elasticities, there was a fall from around 0.75% before the shift to around 0.66% in the low-inflation regime (see Table 1.8). These results are in line with Taylor's hypothesis, that is, the responsiveness of prices to exchange rate fluctuations tend to decline in a low and more stable inflation environment. Likewise, empirically, there was strong evidence in this direction. Comparing our results to those of other studies, in their sample of 11 industrialized countries, BAILLIU and FUJII (2004) found that ERPT in the short-run declined from 0.86% to around 0.71% following a change in the inflation environment. It is important to mention that results from our dynamic panel model corroborate those in the Section 6 regarding the role of inflation levels in explaining the ERPT.

9. Sectoral analysis of ERPT

In the previous sections, we have focused on the overall effect of exchange rate changes into aggregate import prices data rather than on particular industries or products. Thus, in this section, we aim to provide a micro-level analysis of the ERPT as a complement to the previous macro-level evidence. As is well-known, there is a substantial debate about the prevalence of micro or macro factors in explaining the ERPT. A prominent study frequently cited in this regard is CAMPA and GOLDBERG (2005) who differentiated micro-economic from macro-economic explanations for the recent decline in the responsiveness of import prices to exchange rate movements. The authors concluded that changes in the composition of imports toward goods whose prices are less sensitive to exchange rate movements, such as differentiated goods in the manufacturing sector, has been the primary driver behind recent ERPT changes among several OECD countries. Known as "Campa-Goldberg compositional-trade hypothesis",

this phenomenon is considered to explain the lion's share of the decline in pass-through over the past decades (see e.g. GOLDBERG and TILLE, 2008).

Nevertheless, in our study, we have found a substantial role for macroeconomic factors, mainly the inflation regime. Additionally, we check for the importance of the microeconomic factors, mainly the product composition of trade, using disaggregated import price data for each EA country. The methodology for estimation draws heavily on CAMPA, GOLDBERG, and GONZÁLEZ-MÍNGUEZ (2005) and CAMPA and GONZÁLEZ (2006), whose results are updated here. Thus, we reestimate our ERPT equation (1.12) for the disaggregated import price data for all our EA countries except for Luxembourg due to the lack of sufficiently disaggregated data.³⁶ The same industry classification for all the EA countries in the sample is used in order to maintain some comparability of the industry estimations across countries. Then, the disaggregated import price data for each country corresponds to the 1-digit level of disaggregation in the Standard International Trade Classification (SITC) for 9 different industry categories.³⁷ The product disaggregation is as follows: 0. Food and live animals, 1. Beverages and tobacco, 2. Crude materials, inedible, 3. Mineral fuels, 4. Oils, fats and waxes, 5. Chemical products, 6. Basic manufactures, 7. Machines and transport equipment, 8. Miscellaneous manufactured goods (see Table 1.9 in Appendix A.8).³⁸ Because of data availability, the pass-through equation is estimated using monthly data over 1995:01-2010:12. The disaggregated monthly import price data for our 11 EA countries are monthly indexes of import unit values obtained from the database COMEXT of Eurostat.³⁹

When we look to the product composition of imports as reported in Figure 1.10, we detect that the importance of sectors varies across EA countries. on one hand, Belgium,

³⁶We can reestimate the dynamic version of the ERPT equation (1.14) instead of the static one, this gives very similar results for the short-run effect of the exchange rate movements.

³⁷In some empirical studies, higher level of disaggregation are used with 2- and 3-digit SITC product grouping (see e.g. COUGHLIN and POLLARD, 2004; OLIVEI, 2002; YANG, 2007, among others).

³⁸There are no data for category 9 (goods considered as "n.e.s." or not elsewhere specified), which has a residual nature.

³⁹We acknowledge the drawback of using index based on unit values rather than prices which may be problematic for the comparability of goods over time. As explained in the literature, the unit value measures do not properly account either for changes in the definition of product categories over time or for changes in relative demand of similar goods. It is still an aggregate price index, comprising all imported goods in the country within that product category.

Germany and Ireland have the largest share of the manufacturing sectors (SITC 5, 6, 7 and 8) among our sample. For example, about 75% of imports are manufactured products in Germany. On the other hand, Spain, Greece and Portugal corresponds to the countries with the highest portion of mineral fuels (or energy) sector (SITC 3) which contains petroleum products. In Spain, more than 30% of imports are products stemming from energy sector (SITC 3). Besides, within each country, the distribution of imports also varies widely across different product categories.

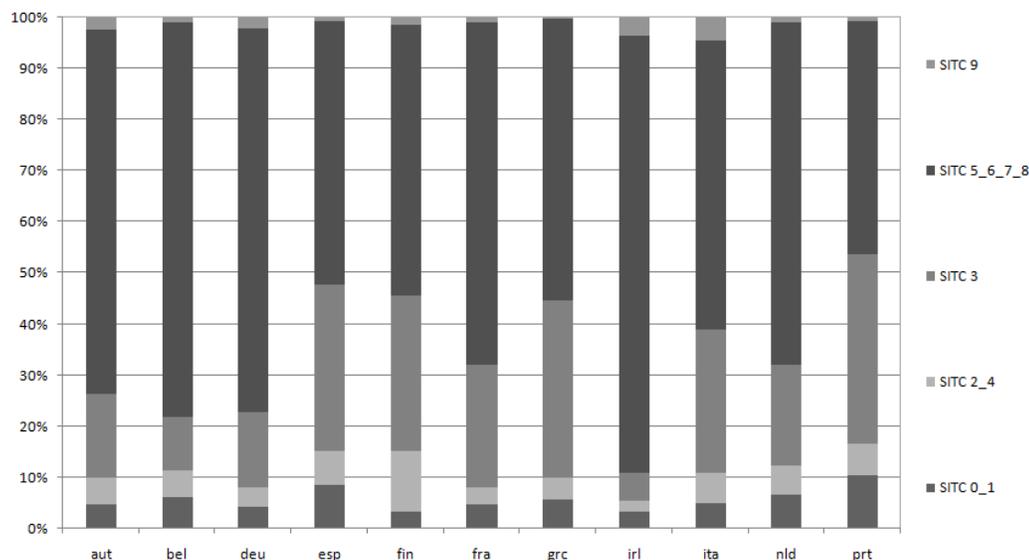
Table 1.9: Standard International Trade Classification

SITC	Industry
SITC 0	Food and live animals
SITC 1	Beverages and tobacco
SITC 0_1	Food, Beverages and tobacco
SITC 2	Crude materials, inedible, except fuels
SITC 3	Mineral fuels, lubricants and related materials
SITC 33	Petroleum, petroleum products and related materials
SITC 4	Animal and vegetable oils, fats and waxes
SITC 2_4	Raw Materials
SITC 5	Chemicals and related products, n.e.s.
SITC 6	Manufactured goods classified chiefly by material
SITC 7	Machinery and transport equipment
SITC 8	Miscellaneous manufactured articles
SITC 5_6_7_8	Manufacturing
SITC 9	Commodities and transactions not classified elsewhere in the SITC

Source: United Nations Statistics Division.

A cursory look of Figure 1.10 shows that the manufacturing sectors (SITC 5, 6, 7 and 8) account for the highest share of imports followed by mineral fuels sector across all EA countries. It is well known, that partial pass-through is a common phenomenon particularly among heterogeneous products (such as manufactured products), while more homogeneous products (such as raw materials) have higher degree of exchange rate transmission (see e.g. CAMPA and GOLDBERG, 2002).⁴⁰ Thus, the divergences in trade composition, as noted in Figure 1.10, would have important implications, especially, in explaining the significant differences in the aggregate import-price pass-through across our EA countries.

⁴⁰As predicted by the law of one price, homogeneous goods must sell for the same price when their prices are converted to a common currency, regardless of where those goods are sold.

Figure 1.10: Share of imports per industry (average over 1995-2010)

Source : COMEXT database of Eurostat.

For purposes of illustration, we plot the correlation between the ERPT elasticities (as computed in Section 6) and the share of different sectors (as percentage of total imports). As reported in Figure A.5 in Appendix A.8, there is a negative relationship between the extent of pass-through and the share of manufacturing sectors (SITC 5, 6, 7 and 8) in total of imports (see upper left subfigure in Figure A.5 in Appendix A.8). This implies that the larger share of differentiated goods (such as manufactured products) in total imports, the lower will be the degree of ERPT into import prices. This negative statistical correlation is more apparent with machinery and transport equipment sector (SITC 7). For a given economy, the larger portion of imported goods stemming from machinery and transport equipment sector, the less would be the extent pass-through (see upper right subfigure in Figure A.5). For the homogeneous goods belonging to energy sector (SITC 3) or raw materials sector (SITC 2 and 4), the link is rather positive with the transmission of exchange rate changes.⁴¹ According to the bottom right subfigure in Figure A.5, the higher are the raw material imports, the more movements in exchange rates are transmitted to import prices. The same positive relationship is found with the energy sector (Mineral fuels, lubricants and related materials).

⁴¹Products belonging to energy and raw materials sectors can be viewed as being closer to classification as imported intermediate goods than food and manufacturing products.

Table 1.10: Sectoral ERPT estimates

Industry	Austria	Belgium	Germany	Spain	Finland	France	Greece	Ireland	Italy	Netherlands	Portugal
SITC 0	0,418 (0,105) [0,024]	0,349 (0,094) [0,002]	0,232 (0,018) [0,000]	0,381 (0,047) [0,001]	0,381 (0,000) [0,000]	0,476 (0,091) [0,063]	0,075 (0,891) -	0,492 (0,065) [0,057]	0,641 (0,054) [0,282]	0,413 (0,230) [0,088]	0,012 (0,622) [0,000]
SITC 1	0,671 (0,385) -	0,786 (0,079) [0,633]	-0,092 (0,196) -	0,800 (0,111) [0,690]	0,409 (0,177) -	-0,107 (0,875) -	0,706 (0,062) [0,437]	1,388 (0,627) -	-0,604 (0,469) -	-0,184 (0,717) -	0,911 (0,031) [0,833]
SITC 2	0,766 (0,079) [0,591]	0,958 (0,020) [0,919]	0,735 (0,203) -	0,728 (0,001) [0,227]	1,175 (0,021) [0,732]	1,006 (0,014) [0,787]	0,960 (0,035) [0,931]	2,415 (0,165) [0,416]	0,791 (0,038) [0,583]	0,986 (0,101) [0,981]	1,076 (0,042) [0,886]
SITC 3	0,938 (0,029) [0,884]	1,014 (0,012) [0,973]	1,048 (0,002) [0,890]	1,186 (0,035) [0,741]	1,222 (0,030) [0,694]	0,828 (0,007) [0,573]	2,760 (0,081) [0,266]	0,932 (0,011) [0,853]	0,865 (0,086) [0,788]	0,843 (0,041) [0,703]	1,534 (0,070) [0,529]
SITC 4	0,838 (0,009) [0,613]	0,809 (0,002) [0,456]	0,790 (0,151) [0,703]	1,071 (0,235) -	0,740 (0,336) [0,736]	1,370 (0,060) [0,612]	1,137 (0,022) [0,783]	0,323 (0,103) [0,001]	0,976 (0,163) [0,973]	1,543 (0,091) [0,551]	2,060 (0,182) -
SITC 5	0,032 (0,506) -	0,438 (0,100) [0,035]	0,023 (0,746) -	0,794 (0,017) [0,535]	0,336 (0,132) [0,003]	-0,092 (0,722) -	-0,084 (0,532) -	-0,326 (0,470) [0,003]	1,003 (0,064) [0,995]	0,913 (0,001) [0,760]	-0,005 (0,945) -
SITC 6	0,200 (0,110) [0,000]	0,515 (0,000) [0,000]	0,262 (0,001) [0,000]	0,694 (0,062) [0,410]	0,427 (0,037) [0,005]	0,685 (0,028) [0,312]	0,762 (0,052) [0,543]	-0,148 (0,727) -	1,000 (0,005) [0,999]	0,751 (0,085) [0,568]	1,635 (0,073) [0,487]
SITC 7	0,005 (0,969) -	0,119 (0,049) [0,000]	0,157 (0,389) -	0,438 (0,091) [0,030]	0,281 (0,526) -	-0,039 (0,601) [0,000]	-0,366 (0,646) [0,087]	1,159 (0,070) [0,804]	0,746 (0,097) [0,573]	0,782 (0,481) -	0,438 (0,395) -
SITC 8	0,174 (0,064) [0,000]	0,233 (0,001) [0,000]	0,283 (0,138) [0,000]	0,694 (0,025) [0,324]	0,609 (0,070) [0,244]	0,518 (0,035) [0,049]	1,074 (0,055) [0,894]	-0,013 (0,972) -	0,453 (0,002) [0,000]	0,986 (0,102) [0,982]	0,833 (0,099) [0,740]

Note: Estimations are based on equation (1.12) using disaggregated import price data. Numbers in parentheses are p -values. For the exchange rate coefficient, p -values in parentheses are based on the null hypothesis of zero ERPT, i.e. $H_0 : \beta_1 = 0$, while p -values in square brackets corresponds to the null of full ERPT, i.e. $H_0 : \beta_1 = 1$.

To provide insights on the “Campa-Goldberg compositional-trade hypothesis”, sectoral ERPT estimates obtained from equation (1.12) using disaggregated import-price data are displayed in Table 1.10. The first remark is the higher variability of the extent of pass-through across and within EA countries. A careful inspection of Table 1.10 reveals that pass-through estimates are usually less than full across industries. The exceptions are mineral fuels (SITC 3) and raw materials (SITC 2 and 4) sectors where the hypothesis of complete ERPT ($H_0 : \beta_1 = 1$) is accepted in most of cases. More precisely, the hypothesis of null ERPT ($H_0 : \beta_1 = 0$) is accepted for Machinery and transport equipment sector (SITC 7) in most of EA countries, except for Belgium, Spain and Ireland. This behavior seems to be present usually in differentiated products. At the other extreme, the mineral fuels (SITC 3) exhibit a full ERPT throughout our country sample, the null of $H_0 : \beta_1 = 1$ is not rejected in any case. The goods included in SITC 3 (such as oil) are examples of relatively homogeneous products. Thus, we can say that our results confirm the conventional wisdom that the more a product is differentiated, the weaker will be the impact of the exchange rate on its import price.

For a better understanding of sectoral pass-through estimates, we summarize results in a more tractable way by reporting summary statistics by industry and by country as in Table 1.11. The most striking result is that ERPT is complete in energy sector (SITC 3) for 100% of our EA countries. Likewise, the responsiveness of import prices in crude materials (SITC 2) is found to be high, with 84% of our country sample having full transmission for this kind of goods. Besides, the large portion of case of zero ERPT is present in beverages and tobacco (SITC 2) and machinery and transport equipment (SITC 7) sectors. The hypothesis of null ERPT is accepted in 64% of cases. Finally, the hypothesis of partial ERPT, i.e. where both of null and full ERPT hypotheses are rejected, is more frequent in food (SITC 0) and miscellaneous manufactured goods (SITC 8). Overall, these results confirm the heterogeneity of the transmission of exchange rate. We found much higher degree of pass-through for more homogeneous goods and commodities, such as oil and raw materials, than for highly differentiated manufactured products, such as machinery and transport equipment. This outcome has an important implication on the evolution of the degree of ERPT over time. As suggested by CAMPA and GOLDBERG (2005), the shift in the composition of imports

toward sectors with lower degrees of pass-through, namely manufactured differentiated goods, would explain the observed decline in ERPT across industrialized countries.

Table 1.11: Summary of sectoral ERPT by industry and country

Industry	Percentage of countries with		
	Full ERPT	Zero ERPT	Partial ERPT
0: Food and live animals	9%	27%	64%
1: Beverages and tobacco	36%	64%	0%
2: Crude materials	82%	18%	0%
3: Mineral fuels	100%	0%	0%
4: Oils, fats and waxes	64%	27%	9%
5: Chemicals products	27%	55%	18%
6: Manufactured goods	55%	9%	36%
7: Machinery and transport equipment	18%	64%	18%
8: Miscellaneous manufactured goods	45%	9%	45%

Country	Percentage of industries with		
	Full ERPT	Zero ERPT	Partial ERPT
Austria	33%	33%	33%
Belgium	44%	0%	56%
Germany	44%	22%	33%
Spain	67%	11%	22%
Finland	44%	33%	33%
France	44%	33%	22%
Greece	67%	33%	0%
Ireland	44%	44%	22%
Italy	78%	11%	11%
Netherlands	56%	44%	0%
Portugal	56%	44%	0%

Note: Full ERPT is the acceptance of $H_0: \beta_1 = 1$, zero ERPT is the acceptance of $H_0: \beta_1 = 0$ and partial ERPT is the reject of both full and zero ERPT hypotheses.

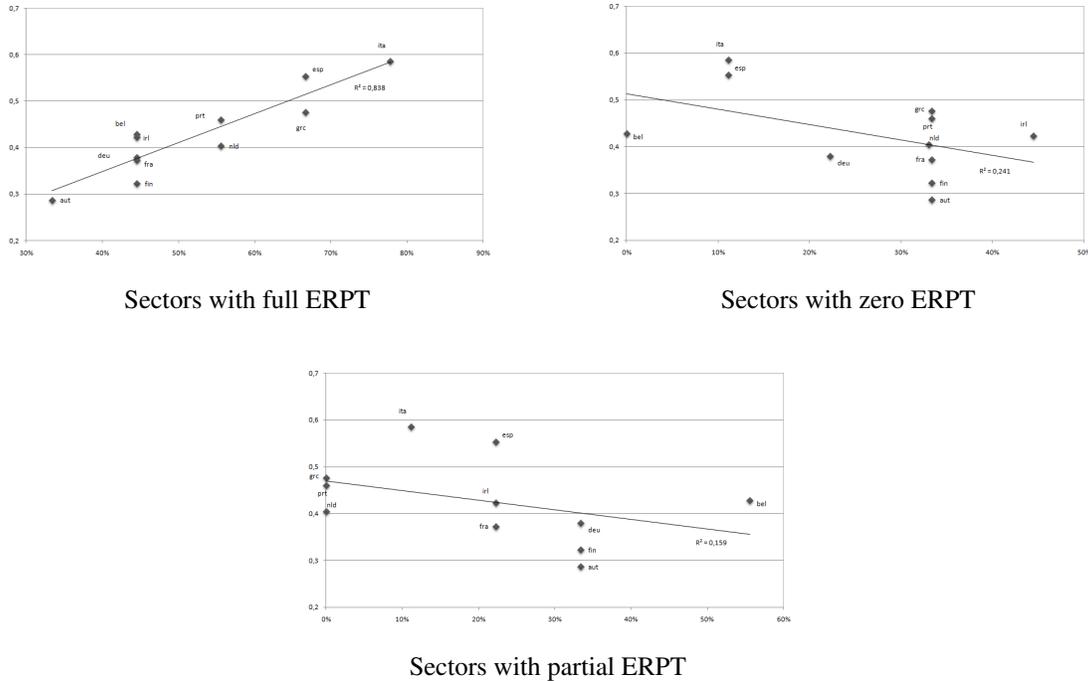
Next, we focus on the sectoral ERPT distribution by country as showed in the bottom of Table 1.11. Results indicate that Spain, Greece and Italy have the largest portion of industries with full ERPT. This group of countries has also a small share of sectors with partial ERPT. For instance, the hypothesis of full ERPT is accepted for 75% of Italian sectors, while in very few cases (11%) the hypothesis of null ERPT is accepted. By contrast, for the Austrian economy, we note a similar distribution of the degree of pass-through, i.e. the same percentage of 33% is found for the respective hypotheses of full, null and partial ERPT. As discussed, the responsiveness of imports prices at the

industry level would explain the observed pass-through at the aggregate level for a given country. Thus, it is interesting to realize that countries with large share of industries having full ERPT, coincide with economies having higher overall rate of pass-through as suggested in Section 6, which is the case of Spain, Greece and Italy. We can say that differences in pass-through at the aggregate level are related to the composition of country imports. The higher share of sectors with lower degrees of pass-through, such as manufacturing sectors, the less the aggregate ERPT will be (and vice versa). Thus, the divergences in the product composition of imports, as in Figure 1.10, can account for a significant amount of the aggregate differences of import price pass-through across countries.

Finally, to clarify the picture, we plot the correlation between the aggregate ERPT, as computed in Section 6, with the share of sectors having full, null and partial ERPT, respectively. As expected, first plot in Figure 1.11 reveals a positive correlation between the overall pass-through and the proportion of industries with full ERPT. The larger share of sectors with complete transmission of exchange rate, the higher would be the response of the aggregate import prices. On the other hand, we find that aggregate pass-through is negatively correlated to the percentage of sectors with null or partial ERPT. We are expecting this result since the more important share of industries having null or full pass-through, the weaker movements in exchange rates are transmitted to import prices.

It appears that the relative importance of different sectors in total import volumes account considerably in the overall observed pass-through rate of a given economy. Generally speaking, our results provides clear support for the “Campa-Goldberg compositional-trade hypothesis”, i.e. the composition of country imports would determine the aggregate response of imports prices to exchange rate movements, and thus, possible differences in overall pass-through rates are due to an heterogeneous industry composition of trade across countries. Nevertheless, it is worth stressing that, contrary to CAMPA and GOLDBERG (2005), in our study we found a substantial role for some macroeconomic factors, mainly inflation environment (see Section 6 and Section 8). While CAMPA and GOLDBERG (2005) found that the composition of industries in a country’s import basket is by far the primary driver the behavior of pass-through into import prices.

Figure 1.11: Correlation aggregate ERPT and percentage of industries with full, null and partial ERPT



Note: y-axis: ERPT to import prices estimated from equation (1.12) over 1990-2010; x-axis: share of sectors with full, null and partial ERPT.

10. Conclusion

In this chapter an update of the ERPT estimates is provided for 12 EA countries. First, using quarterly data over the period of 1990-2010, we don't find a wide heterogeneity in the degree of pass-through across the 12 EA countries, in contrast to previous empirical works. This is not surprising since previous studies used too few observations for the EA era, while in our work, the time span for the analysis of the post-EA era is rather long (until the end of 2010). Since the process of monetary union has entailed some convergence towards more stable macroeconomic conditions, it is expected to find a relative low and less dispersed ERPT across EA Member States.

Concerning the macro determinants, we found a positive relationship between ERPT and inflation in line with TAYLOR (2000), while no significant role for the degree of openness, measured as the ratio of imports to GDP. Assessing the stability of pass-through elasticities, we find very weak evidence of decline around 1999. However, our results reveal that the pass-through estimates appears to trend down since the beginning of the 1990s. We notice that the observed decline was synchronous to the shift towards reduced inflation regime in our sample of 12 EA countries. It is interesting to note that when we estimate our pass-through equation over 1979:2-1990:2, we point out more pronounced cross-differences in ERPT than recorded over 1990:3-2010:4. There was divergent macroeconomic conditions across EA countries during the 1980s, especially between peripheral and core economies. Thereby, it is expected that the general process of European convergence, which has started before the introduction of the Euro in 1999, would entail a reduction in the variability of pass-through within EA members states.

Thereafter, within the dynamic panel data framework, we confirm the non-significant decline of the import-price sensitivity to exchange rate since the formation of the euro. However, the important role played by inflation environment was confirmed once again. We found that the responsiveness of prices to exchange rate fluctuations tend to decline in a low and more stable inflation environment. Moreover, our findings suggest that the weakness of the euro during the first three years of the monetary union has raised significantly the extent of pass-through. We pretend that this outcome would explain why the sensitivity of the import prices did not fall since 1999. Another important which is the ERM crisis over 1992-1993 seems to do not affect the sensitivity of import prices to exchange rate movements in the EA.

Finally, using disaggregated import prices data, it appears that the product composition of imports would determine the aggregate ERPT of an economy, and thus, cross-country differences in pass-through rates may be due to an heterogeneous industry composition of trade across countries.

Appendix A

A.1. Deriving the ERPT elasticity

The profit maximization problem yields the following second order conditions:

$$\begin{vmatrix} \frac{\delta^2 \Pi}{\delta^2 p^x} & \frac{\delta^2 \Pi}{\delta p^x \delta p^m} \\ \frac{\delta^2 \Pi}{\delta p^m \delta p^x} & \frac{\delta^2 \Pi}{\delta^2 p^m} \end{vmatrix} > 0, \frac{\delta^2 \Pi}{\delta^2 p^x} < 0, \text{ and } \frac{\delta^2 \Pi}{\delta^2 p^m} < 0. \quad (\text{A.1})$$

According to the second inequality in (A.1):

$$\frac{\delta^2 \Pi}{\delta^2 p^x} = \frac{\delta^2 q^x}{\delta^2 p^x} \left(\frac{p^x}{\mu^x} - w^* \phi' \right) + \frac{\delta q^x}{\delta p^x} \left(\frac{1}{\mu^x} (1 - \eta^{\mu^x}) - w^* \phi'' \frac{\delta q^x}{\delta p^x} \right) < 0 \quad (\text{A.2})$$

where $\eta^{\mu^x} = \frac{\delta \mu^x}{\delta p^x} \frac{p^x}{\mu^x} \leq 0$, is the elasticity of the markup with respect to the price in foreign country.

By the first order condition (1.8), $\left[\frac{p^x}{\mu^x} - w^* \phi' \right] = 0$, the sign of (A.2) depends on the sign of $\frac{\delta q^x}{\delta p^x} \left(\frac{1}{\mu^x} (1 - \eta^{\mu x}) - w^* \phi'' \frac{\delta x^x}{\delta p^x} \right)$. Assuming that demand is well behaved, $\frac{\delta q^x}{\delta p^x} < 0$, thus, $\left(\frac{1}{\mu^x} (1 - \eta^{\mu x}) - w^* \phi'' \frac{\delta x^x}{\delta p^x} \right) > 0$.

Similarly, the third inequality in (A.1) is as follow:

$$\frac{\delta^2 \Pi}{\delta^2 p^m} = \frac{\delta^2 q^m}{\delta^2 p^m} \left(\frac{e^{-1} p^m}{\mu^m} - w^* \phi' \right) + \frac{\delta x^m}{\delta p^m} \left(\frac{e^{-1}}{\mu^m} (1 - \eta^{\mu m}) - w^* \phi'' \frac{\delta q^m}{\delta p^m} \right) < 0 \quad (A.3)$$

where $\eta^{\mu m} = \frac{\delta \mu^m}{\delta p^m} \frac{p^m}{\mu^m} \leq 0$, is the elasticity of the markup with respect to the price in the importing country.

By the first order condition (1.9), $\left[\frac{e^{-1} p^m}{\mu^m} - w^* \phi' \right] = 0$, the sign of (A.3) will depend on the sign of $\frac{\delta q^m}{\delta p^m} \left(\frac{e^{-1}}{\mu^m} (1 - \eta^{\mu m}) - w^* \phi'' \frac{\delta x^m}{\delta p^m} \right) < 0$. If demand is well behaved, $\frac{\delta q^x}{\delta p^x} < 0$, consequently, $\left(\frac{e^{-1}}{\mu^m} (1 - \eta^{\mu m}) - w^* \phi'' \frac{\delta x^m}{\delta p^m} \right) > 0$.

The response of import price p^m with respect to a change in the exchange rate is obtained by using the implicit function theorem to the first-order condition given in the text (equations (1.8) and (1.9):

$$\frac{\delta p^m}{\delta e} = \frac{\frac{e \cdot w^* \phi'}{\mu^x} (1 - \eta^{\mu x}) (1 + \eta^{w^* e}) - w^* \phi'' \frac{\delta q^x}{\delta p^x} \frac{p^m}{\mu^m}}{\frac{e(1 - \eta^{\mu x})(1 - \eta^{\mu m})}{\mu^x \mu^m} - w^* \phi'' \left[\frac{(1 - \eta^{\mu x})}{\mu^m} \frac{\delta q^x}{\delta p^x} + \frac{e(1 - \eta^{\mu m})}{\mu^m} \frac{\delta q^x}{\delta p^x} \right]} \quad (A.4)$$

where $\eta^{w^* e} = \frac{\delta w^*}{\delta e} \frac{e}{w^*}$ is the elasticity of price factors with respect to the exchange rate.

Supposing that marginal costs are constant, $w^*\phi'' = 0$, equation (A.4) becomes:

$$\frac{\delta p^m}{\delta e} = \frac{e \cdot w^* \phi' (1 + \eta^{w^*e})}{\frac{e}{\mu^m} (1 - \eta^{\mu m})} \quad (\text{A.5})$$

According to (1.9):

$$\frac{p^m}{\mu^m} = e \cdot w^* \phi' \quad (\text{A.6})$$

Therefore:

$$\frac{\delta p^m}{\delta e} = \frac{\frac{p^m}{\mu^m} (1 + \eta^{w^*e})}{\frac{e}{\mu^m} (1 - \eta^{\mu m})} = \frac{p^m (1 + \eta^{w^*e})}{e (1 - \eta^{\mu m})} > 0 \quad (\text{A.7})$$

Rearranging the latter equation gives us the ERPT elasticity:

$$\text{ERPT} = \frac{\delta p^m}{\delta e} \frac{e}{p^m} = \frac{1 + \eta^{w^*e}}{1 - \eta^{\mu m}} \geq 0 \quad (\text{A.8})$$

A.2. Stationary Tests

Table A.1: Unit Root Tests for main series

	Austria	Belgium	Germany	Spain	Finland	France
Δmp_t	-3,945**	-5,603**	-7.159**	-5.344**	-8.064**	-6,824**
	-4,857*	-6,169**	-5.279*	-5.628**	-8.749**	-7,096**
Δe_t	-7,069**	-6.975	-7.071**	-6.953**	-6.580**	-6,989**
	-7,615**	-7.499**	-7.619**	-7.491**	-6.939**	-7,509**
Δw_t^*	-6.709**	-7.033**	-7.212**	-7.323**	-5.682**	-6.708**
	-7.311**	-7.889**	-7.865**	-5.048*	-4.338	-7.341**
gap_t	-4,680**	-4.861**	-4.461**	-4.084**	-3.470*	-3,964**
	-3,734	-5.127*	-8.177**	-4.746	-4.498	-4,226
Δppi_t	-6,212**	-3.400*	-4.116**	-3.932**	-4.718**	-4,254**
	-4,695	-4.252	-4.316	-3.805	-3.927	-4,302*
Δoil_t	-8,353**	-8.353	-8.353**	-8.353**	-8.353**	-8,673**
	-8.943**	-8.943**	-8.943**	-8.943**	-8.943**	-8,943**
Δgdp_t	-4,243**	-5.164**	-8.732**	-3.067*	-5.494**	-4,692**
	-3,587	-5.572**	-9.275**	-4.286	-6.595**	-4,089
	Greece	Ireland	Italy	Luxembourg	Netherlands	Portugal
Δmp_t	-6.053**	-8.056**	-5.587	-8.475**	-6.743**	-5.051**
	-7.494**	-8.756**	-5.980**	-8.950**	-6.291**	-6.239**
Δe_t	-6.576**	-6.898**	-7.044**	-6.975**	-7.057**	-6.656**
	-7.295**	-7.410**	-7.855**	-7.499**	-7.610**	-7.162**
Δw_t^*	-6.281**	-6.903**	-6.414**	-6.615**	-6.947**	-6.289**
	-4.669	-7.541**	-7.989**	-7.297**	-7.759**	-5.16994*
gap_t	-4.180**	-3.021*	-3.965**	-3.366*	-3.015*	-2.645*
	-4.646	-4.450	-4.346	-4.375	-3.614	-4.131
Δppi_t	-4.231**	-6.810	-8.353	-4.560**	-2.960*	-4.430**
	-3.012	-7.291**	-8.943**	-4.813*	-4.213	-5.165*
Δoil_t	-8.353**	-3.076*	-5.025**	-8.353**	-8.353**	-8.357**
	-8.943**	-4.080	-5.603**	-8.943**	-8.943**	-8.943**
Δgdp_t	-6.863**	-8.353	-6.492**	-10.403**	-5.693**	-6.420**
	-8.385**	-8.943**	-7.871**	-6.238**	-6.402**	-4.461

Note: First and second row for each series report ADF and ZIVOT and ANDREWS (1992) test, respectively. **,* denotes rejection the null hypothesis of unit root at 5% and 10%, respectively. ZIVOT and ANDREWS (1992) test allow for one single break under the alternative hypothesis. Lag selection: Akaike (AIC). Maximum lags number = 8.

Table A.2: Cointegration tests for benchmark model

	Austria	Belgium	Germany	Spain	Finland	France
EG test	-2,184	-1,441	-1,877	-1,986	-3,501**	-2,730
GH test						
<i>constant</i>	-4,155	-3,568	-3,892	-3,864	-4,423	-4,243
<i>constant & slope</i>	-5,273	-5,014	-6,620*	-5,572	-6,132	-6,295*
	Greece	Ireland	Italy	Luxembourg	Netherlands	Portugal
EG test	-3,63295**	-3,18447*	-1,574	-2,781	-2,654	-1,658
GH test						
<i>constant</i>	-5,147	-3,728	-5,354*	-3,038	-5,221*	-4,444
<i>constant & slope</i>	-5,937	-3,947	-6,918**	-5,146	-6,887**	-5,601

Note: **, * the null hypothesis of unit root in the residuals (no cointegration) is rejected at 5% and 10%, respectively. First reports ENGLE and GRANGER (1987) test. The following rows correspond to GREGORY and HANSEN (1996) tests. Specifications for GH tests include both a constant and a time trend. Lag selection: Akaike (AIC). Maximum lags number = 8.

A.3. Robustness checks

A.3.1. Robustness check with additional explanatory variables

Table A.3: Estimation results with producer prices

	Austria	Belgium	Germany	Spain	Finland	France
Constant	0,030 (0,000)	-0,007 (0,847)	0,020 (0,685)	-0,002 (0,950)	0,066 (0,494)	-0,001 (0,550)
Δe_t	0,357 (0,000)	0,428 (0,000)	0,406 (0,000)	0,553 (0,000)	0,418 (0,000)	0,376 (0,000)
Δw_t^*	0,565 (0,000)	0,638 (0,000)	0,528 (0,000)	0,664 (0,000)	0,651 (0,000)	0,621 (0,000)
gap_t	0,016 (0,874)	0,311 (0,003)	0,025 (0,457)	0,100 (0,364)	0,063 (0,460)	0,166 (0,082)
Δppi_t	-0,233 (0,166)	0,002 (0,859)	-0,008 (0,633)	0,001 (0,959)	-0,022 (0,485)	-0,428 (0,061)
Observations	81	82	82	82	82	82
R^2	0,898	0,603	0,719	0,590	0,426	0,653
	Greece	Ireland	Italy	Luxembourg	Netherlands	Portugal
Constant	0,068 (0,000)	0,044 (0,707)	0,002 (0,955)	0,049 (0,433)	-0,003 (0,881)	0,019 (0,393)
Δe_t	0,423 (0,000)	0,408 (0,000)	0,590 (0,000)	0,454 (0,000)	0,404 (0,000)	0,462 (0,000)
Δw_t^*	0,699 (0,000)	0,308 (0,004)	0,783 (0,000)	0,668 (0,000)	0,636 (0,000)	0,699 (0,000)
gap_t	0,133 (0,094)	0,149 (0,052)	0,172 (0,051)	-0,082 (0,357)	0,118 (0,037)	0,020 (0,756)
Δppi_t	-0,020 (0,002)	-0,012 (0,746)	0,001 (0,926)	-0,013 (0,539)	0,001 (0,904)	-0,006 (0,429)
Observations	78	81	79	82	82	82
R^2	0,658	0,407	0,799	0,296	0,734	0,652

Note: Estimation are based on equation 1.12 including the producer prices Δppi_t as additional explanatory variable. Numbers in parentheses are p -values.

Table A.4: Estimation results with oil prices

	Austria	Belgium	Germany	Spain	Finland	France
Constant	0,029 (0,000)	-0,001 (0,762)	-0,003 (0,036)	0,000 (0,971)	-0,005 (0,186)	-0,004 (0,042)
Δe_t	0,244 (0,000)	0,445 (0,000)	0,397 (0,000)	0,525 (0,000)	0,291 (0,000)	0,303 (0,000)
Δw_t^*	0,321 (0,006)	0,678 (0,000)	0,513 (0,000)	0,580 (0,000)	0,418 (0,002)	0,451 (0,000)
gap_t	-0,012 (0,905)	0,311 (0,003)	0,030 (0,362)	0,100 (0,358)	0,031 (0,697)	0,088 (0,280)
Δoil_t	0,017 (0,185)	-0,006 (0,585)	0,012 (0,189)	0,016 (0,232)	0,018 (0,289)	0,029 (0,000)
Observations	81	82	82	82	82	82
R^2	0,898	0,604	0,711	0,598	0,331	0,708
	Greece	Ireland	Italy	Luxembourg	Netherlands	Portugal
Constant	0,007 (0,009)	0,006 (0,092)	0,005 (0,019)	0,011 (0,009)	0,000 (0,913)	0,002 (0,341)
Δe_t	0,501 (0,000)	0,404 (0,000)	0,578 (0,000)	0,424 (0,000)	0,355 (0,000)	0,432 (0,000)
Δw_t^*	0,853 (0,000)	0,298 (0,029)	0,745 (0,000)	0,595 (0,001)	0,516 (0,000)	0,613 (0,000)
gap_t	0,056 (0,527)	0,149 (0,053)	0,179 (0,039)	-0,085 (0,342)	0,118 (0,027)	0,008 (0,900)
Δoil_t	-0,021 (0,086)	0,002 (0,883)	0,005 (0,617)	0,010 (0,610)	0,020 (0,005)	0,015 (0,121)
Observations	78	81	82	82	82	82
R^2	0,594	0,406	0,794	0,295	0,761	0,661

Note: Estimation are based on equation 1.12 including the oil prices Δoil_t in US dollar as additional explanatory variable. Numbers in parentheses are p -values.

A.3.2. Robustness check with alternative proxy for demand conditions

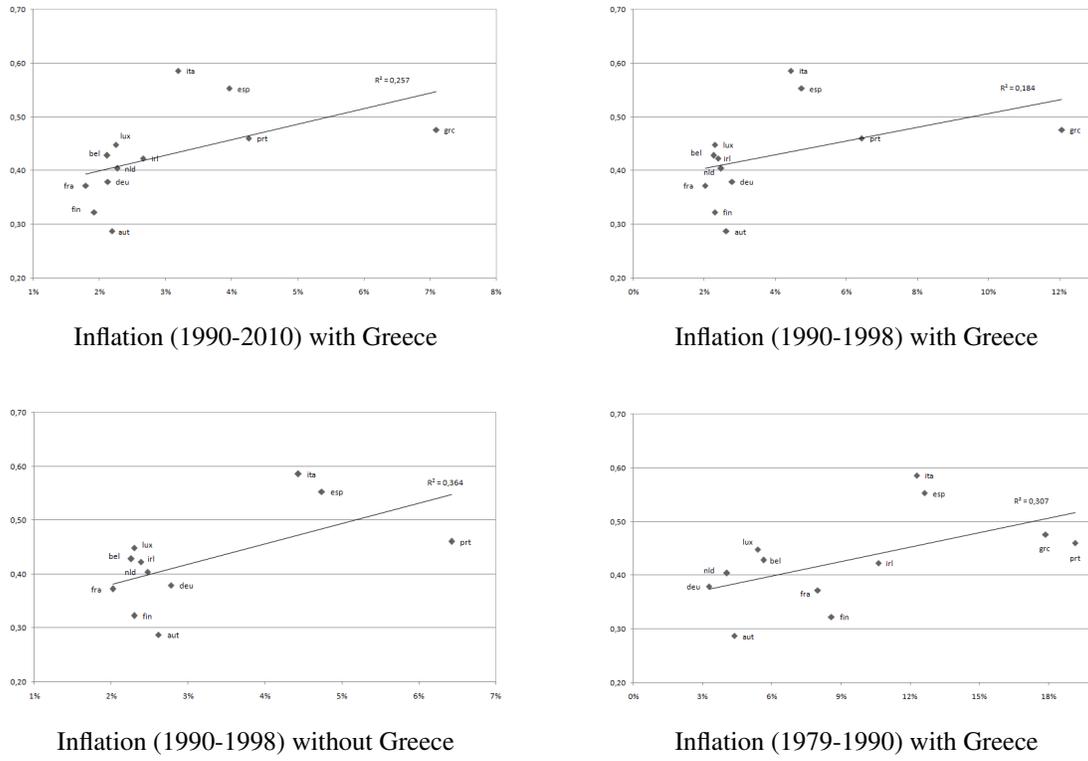
Table A.5: Estimation results with real GDP (growth rate)

	Austria	Belgium	Germany	Spain	Finland	France
Constant	0,028 (0,000)	-0,003 (0,290)	0,010 (0,838)	-0,001 (0,734)	-0,007 (0,088)	-0,005 (0,010)
Δe_t	0,277 (0,000)	0,399 (0,000)	0,411 (0,000)	0,547 (0,000)	0,320 (0,000)	0,363 (0,000)
Δw_t^*	0,410 (0,000)	0,550 (0,000)	0,481 (0,000)	0,654 (0,000)	0,503 (0,000)	0,610 (0,000)
Δgdp_t	0,100 (0,672)	0,572 (0,035)	0,138 (0,222)	0,114 (0,559)	0,107 (0,470)	0,172 (0,392)
Observations	81	82	82	82	82	82
R^2	0,896	0,539	0,715	0,594	0,334	0,651
	Greece	Ireland	Italy	Luxembourg	Netherlands	Portugal
Constant	0,009 (0,001)	0,004 (0,253)	0,004 (0,052)	0,010 (0,021)	-0,001 (0,383)	0,000 (0,900)
Δe_t	0,477 (0,000)	0,406 (0,000)	0,568 (0,000)	0,421 (0,000)	0,389 (0,000)	0,449 (0,000)
Δw_t^*	0,760 (0,000)	0,310 (0,005)	0,745 (0,000)	0,621 (0,000)	0,609 (0,000)	0,673 (0,000)
Δgdp_t	-0,068 (0,324)	0,151 (0,165)	0,253 (0,145)	0,102 (0,398)	0,195 (0,113)	0,304 (0,015)
Observations	78	81	82	82	82	82
R^2	0,607	0,391	0,787	0,289	0,727	0,676

Note: Estimation are based on equation 1.12 including the the growth rate of real GDP Δpib_t instead of the output gap as a proxy for the changing in the domestic demand. Numbers in parentheses are p -values.

A.4. The connection between pass-through and rate of inflation

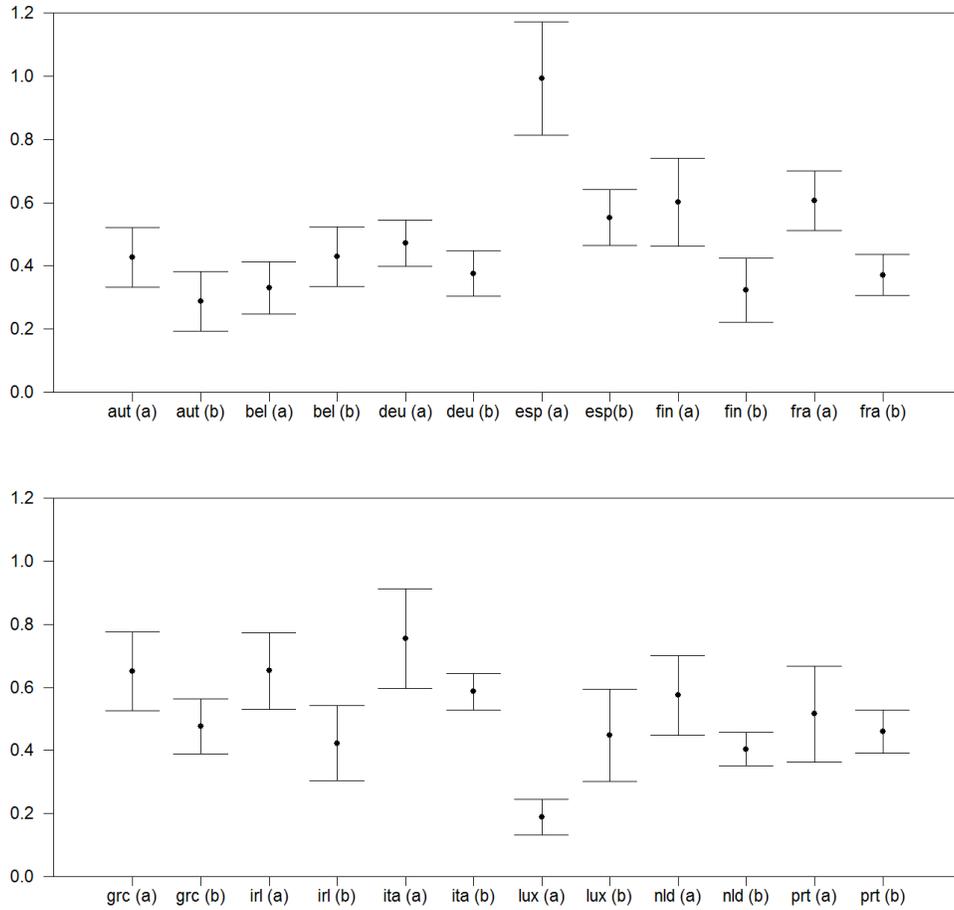
Figure A.1: Correlation between ERPT and different inflation periods



Note: y-axis: ERPT to import prices estimated from equation (1.12) over 1990-2010; x-axis: average of inflation.

A.5. ERPT estimates with 95% confidence intervals

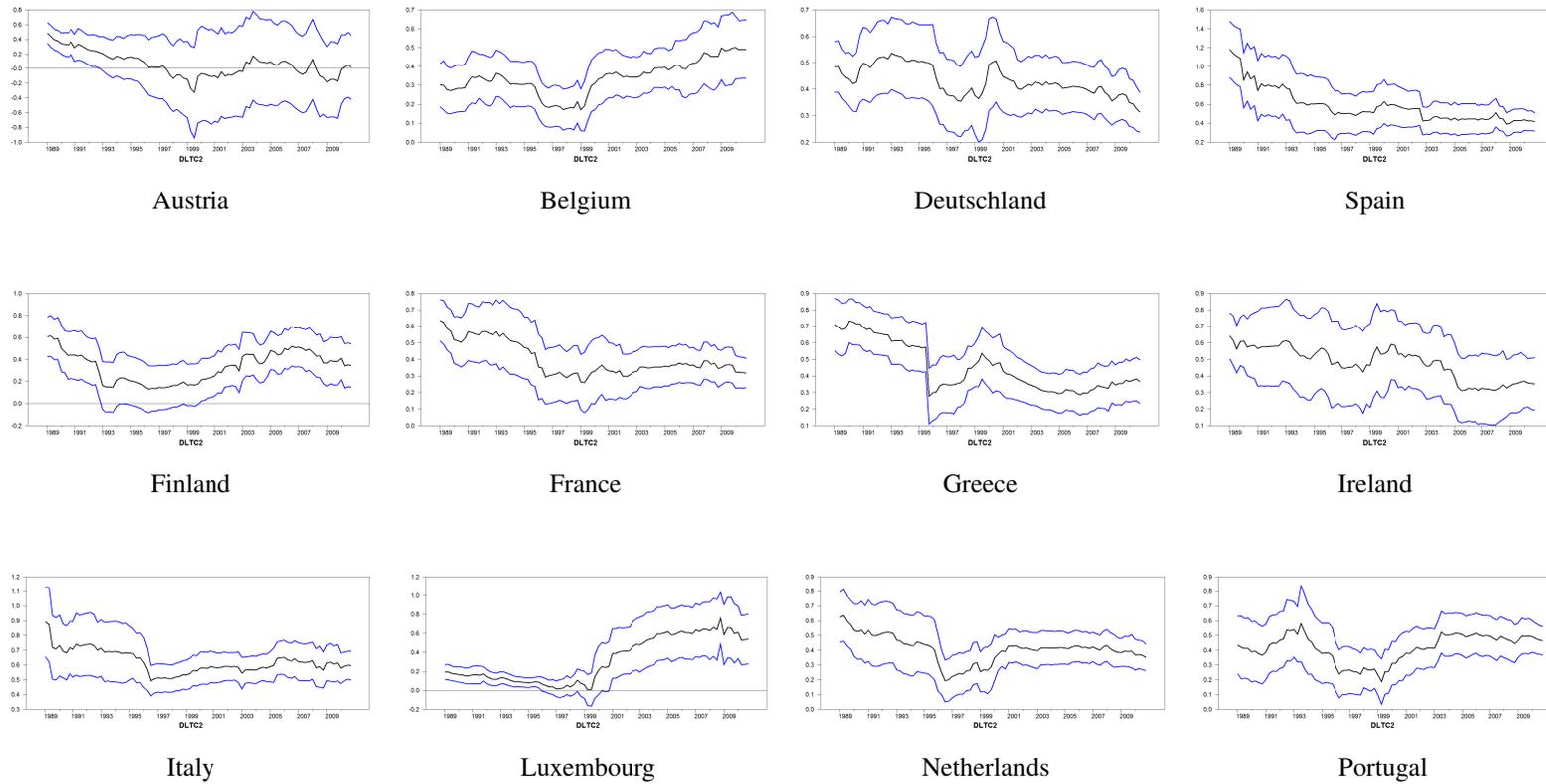
Figure A.2: ERPT point estimates



Note: Figures report ERPT estimates with 95% confidence intervals confidence intervals. (a) corresponds to the ERPT over 1979:2-1990:2, while (b) corresponds to the ERPT over 1990:3-2010:4.

A.6. Moving window estimates with standard error bands

Figure A.3: Moving window ERPT with HAC standard errors



A.7. Identified Structural Breaks in the CPI Inflation Series

Table A.6: BAI and PERRON (2003) multiple break test

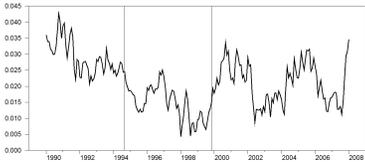
	Austria	Belgium	Germany	Spain	Finland	France	Greece	Ireland	Italy	Luxembourg	Netherlands	Portugal
BIC												
0	-9,191	-9,671	-8,658	-8,363	-8,313	-9,795	-5,594	-8,568	-8,241	-9,474	-9,641	-6,828
1	-9,945	-10,009	-9,605	-9,410	-9,019	-10,334	-7,155	-9,030	-9,983	-9,866	-10,059	-8,312
2	-10,081	-10,112	-9,541	-9,458	-8,960	-10,370	-8,046	-9,227	-10,381	-10,400	-10,230	-8,316
3	-9,967	-10,039	-9,527	-9,881	-8,981	-10,822	-7,979	-9,375	-10,339	-10,372	-10,481	-8,221
LWZ												
0	-9,149	-9,629	-8,615	-8,321	-8,271	-9,752	-5,551	-8,526	-8,198	-9,432	-9,599	-6,786
1	-9,860	-9,924	-9,520	-9,325	-8,934	-10,249	-7,070	-8,946	-9,898	-9,782	-9,974	-8,227
2	-9,911	-9,942	-9,370	-9,288	-8,790	-10,200	-7,876	-9,057	-10,211	-10,230	-10,060	-8,146
3	-9,711	-9,783	-9,270	-9,625	-8,725	-10,566	-7,722	-9,119	-10,083	-10,116	-10,225	-7,965
Sup_F(m)												
1	89,058	34,568	123,119	143,549	81,443	58,156	287,721	48,582	358,811	40,505	43,362	261,028
2	65,730	29,489	65,005	88,477	44,096	38,643	443,772	45,076	315,335	69,427	39,713	147,645
3	42,762	20,962	49,112	114,396	36,113	61,399	303,005	44,852	221,939	51,236	47,108	98,309
Sup_F(m m-1)												
(1 0)	89,058	34,568	123,119	143,549	81,443	58,156	287,721	48,582	358,811	40,505	43,362	261,028
(2 1)	19,922	17,072	3,789	12,392	4,212	11,417	120,324	25,495	46,577	63,350	23,148	8,899
(3 2)	0,219	3,029	7,314	48,075	9,965	51,484	3,404	20,384	5,269	6,267	29,846	1,552
No. of breaks												
BIC	2	2	1	3	1	3	2	3	2	2	3	2
LWZ	2	2	1	3	1	3	2	3	2	2	3	1
Sequential	2	2	1	1	1	1	2	3	2	2	3	1
Break dates	1995:06 2000:01	1994:09 1999:11	1994:06	1990:01	1992:12 1999:10	1992:12 2000:05	1995:01 1999:04	1992:09 1999:12 2003:04	1992:02 1996:08	1994:02 1999:11	1994:12 2000:12 2002:12	1994:06

Note: BIC is the Bayesian Information Criterion (BIC) suggested by Yao (1988) for break selection, while LWZ is a modified Schwarz criterion for break selection proposed by Liu et al. (1997). The $Sup_F(m)$ statistics tests for the null hypothesis of no structural break against m ($m = 1, 2, 3$) breaks. The $Sup_F(m|m-1)$ statistics for the null hypothesis of $m-1$ structural breaks against m ($m = 1, 2, 3$) structural breaks. Last row provided suggested break dates based on the results of these sequential tests ($Sup_F(m)$ and $Sup_F(m|m-1)$).

Figure A.4: Structural Breaks in Inflation series



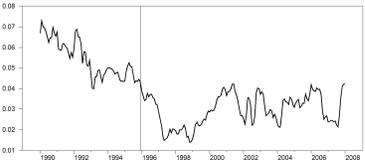
Austria



Belgium



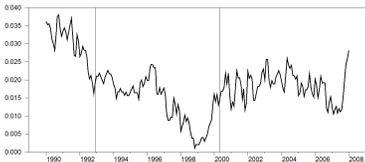
Deutschland



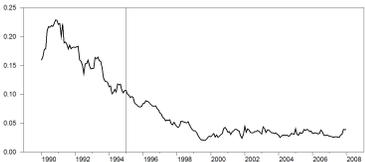
Spain



Finland



France



Greece



Ireland



Italy



Luxembourg



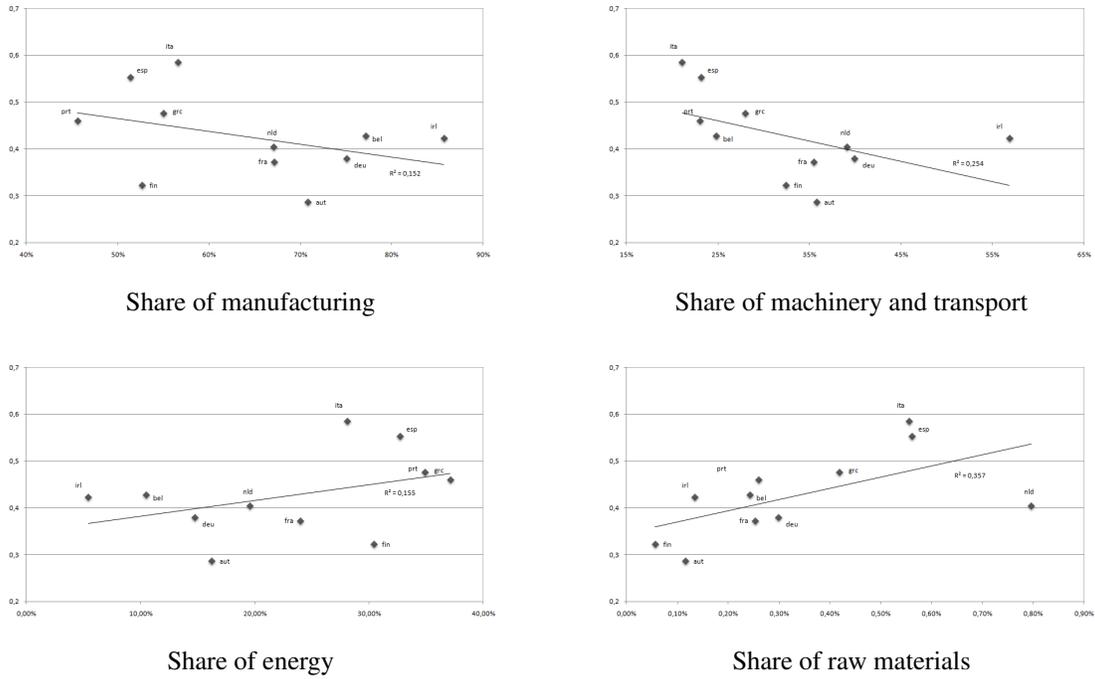
Netherlands



Portugal

A.8. ERPT at the sectoral level

Figure A.5: Correlation between ERPT and share of sector (% of total imports)



Note: y-axis: ERPT to import prices estimated from equation (1.12) over 1990-2010; x-axis: share of sector in total imports (average over 1995-2010).

Chapter 2

Long-run Exchange Rate Pass-through into Import Prices: Evidence from New Panel Data Techniques

1. Introduction

The issue of exchange rate pass-through (ERPT) into domestic prices has long been of interest in debates about the conduct of monetary policy and the choice exchange rate regime. By definition, this concept refers to the degree of sensitivity of import prices to a one percent change in exchange rates in the importing nation's currency. It is commonly argued in pass-through literature that the import prices do not move one-to-one following exchange variations, that is, ERPT is found to be incomplete. Moreover, several industrialized countries have experienced decline in pass-through since the early 1990's. However it is still difficult to answer the question of what factors exactly have caused this trend. In fact, there are several explanations for the reducing pass-through mechanism. From a macroeconomic perspective, the moving towards more stable inflation environment has played an important role in the recent fall in ERPT. This positive correlation between inflation and the degree of pass-through has put forth

by Taylor (2000). Known as Taylor's hypothesis, this argues that countries with low-inflation environment as a result of more credible monetary policies would experience a reduced degree of pass-through. Thus inflation regime can be considered as one of the sources of ERPT differences across countries. For instance, it is arguable that pass-through is always higher in developing economies with more than one-digit level of inflation.

In fact, there are several factors influencing ERPT that are often discussed in pass-through literature. In addition to the inflation environment, CAMPA and GOLDBERG (2005) have tested the importance of other macroeconomic variables that affecting the pass-through, namely, monetary policy stability, country size and exchange rate volatility. The authors found that only exchange rate volatility affects in a statistically significant way the degree of pass-through. In their study, CHOUDHRI and HAKURA (2006) show that ERPT is positively correlated to the average of inflation rate and the inflation and exchange rate volatility, but no significant role for the degree of openness was founded. The present paper follows this strand of literature and, therefore, analyzes the role of some macroeconomic variables that may account for the cross-country differences in pass-through. In a sample of 27 OECD countries, we address the question of whether inflation rate, degree of openness and exchange rate volatility are potential sources of heterogeneity in ERPT. Using panel threshold model introduced by HANSEN (1999), we show that our sample of countries can be classified into different groups according to their macroeconomic regimes. This enables us to test the presence of regime-dependence in ERPT mechanism. To the best of our knowledge, this is the first study that applies panel threshold method in this context.

Another important issue in the literature concerns the long-run equilibrium in the pass-through equation. In fact, several empirical studies have failed to find evidence of cointegrating relationship in the data. As discussed in panel cointegration literature (see e.g. PEDRONI, 1999, 2001, 2004; BREITUNG and PESARAN, 2005, among others), conventional nonstationary tests have low power in small sample sizes, so adding the cross-section dimension to the time series dimension would increase the power of these tests. Therefore, we propose to use panel data cointegrating techniques to restore the long-run equilibrium in ERPT relationship.

THE PURPOSE OF this chapter is three-fold: first, we begin by measuring the long-run ERPT into import prices index for 27 OECD countries. We follow PEDRONI (2001) methodology by applying FMOLS and DOLS group mean estimators. Little is said about long run pass-through in this context, and the aim of our paper is to fill this gap by using these recent panel data techniques. Second, we provide insights into the factors underlying cross-country differences in pass-through elasticities. To this end, we explore three macroeconomic determinants, i.e. inflation rate, degree of openness and exchange rate volatility which are potential sources of heterogeneity in ERPT. Due to the important implications of incomplete pass-through for monetary union, in the final part of our analysis, we focus on the case of the euro area by taking a sub-sample of 12 European countries. Our goal is to assess the behavior of ERPT since the collapse of Breton-Woods era and try to relate it to the change in the inflation environment.

To preview our results, we first provide a strong evidence of incomplete ERPT in our panel 27 OECD countries. On the long run, import prices do not move one-to-one following exchange rate depreciation. Both FM-OLS and DOLS estimators show that pass-through elasticity does not exceed 0.70%. When considering individual estimates for our panel of 27 countries, we can note a cross-country difference in the long run ERPT. Especially, there is an evidence of complete pass-through for 5 out of 27 countries, namely, Czech Republic, Italy, Korea, Luxembourg and Poland. Second, when split our sample in different country regimes, we find that countries with higher inflation regime and more exchange rate volatility would experience a higher degree of pass-through. For the degree of openness, our results provide a weak evidence for a positive link between import share and ERPT. When considering the sub-sample of euro area countries, we find a steady decline in the degree of pass-through throughout the different exchange rate arrangements: ERPT elasticity was close to unity during the *snake-in-the tunnel* period while it is about 0.50% since the formation of the euro area.

The remainder of the **Chapter 2** is organized as follows. Section 2 provides an overview of the literature on ERPT and discusses some macro-determinants that may explain cross-country differences in pass-through. Section 3 describes the analytical framework that underlies our empirical specification and the data used in the study. In Section 4, we discuss the empirical methodology used to test stationarity and cointegration in panel. Results of the empirical analysis for our panel of 27 OECD

countries as well as for each individual are presented in Section 5. Section 6 discusses the main macroeconomic factors determining ERPT. In Section 7, we focus on The EMU countries by assessing how pass-through has changed over time. Section 8 concludes.

2. Overview of the literature

MENON (1995) and GOLDBERG and KNETTER (1997) gave a comprehensive review of a large body of empirical literature which deals with the issue of pass-through to import prices. The main finding of this literature is that import prices do not fully respond to a depreciation or appreciation in the domestic currency. Especially, this finding remains strong in the short run due to the staggered price setting, and pass-through seems to be much lower than in the longer run. However, price adjustment may be incomplete even in the long run, micro-determinants like pricing strategies of firms is one of major reason of partial ERPT. In a seminal papers, DORNBUSCH (1987) and KRUGMAN (1987) justifies incomplete pass-through as a result of firms' markup adjustment depending on market destination. Within imperfect competition market, exporters can practice a pricing-to-market (PTM hereafter) strategy by setting different prices for different destination markets.¹ If the firms keep a constant markup, import prices move one-to-one to changes in exchange rates, and there is no evidence of PTM. This latter case refers to denomination of imports in the currency of the exporting country which is called producer-currency pricing (PCP). And if the firm's markup decreases following destination market currency depreciation, PTM occurs and pass-through to import prices is less than complete. When prices do not to vary in the currency of importing country, this refers to local-currency pricing (LCP) strategy and pass-through would be equal to zero.

In a more recent literature, there has been a growing interest in examining the relationship of ERPT and macroeconomic factors. One of the most convincing factors is the inflation environment in each country. This latter macro-determinant is brought by TAYLOR (2000) who argues that the responsiveness of prices to exchange rate

¹Pricing-to-market is defined as the percent change in prices in the exporter's currency due to a one percent change in the exchange rate. Thus, the greater the degree of pricing-to-market, the lower the extent of exchange rate pass-through.

fluctuations depends positively on inflation. So pass-through tends to increase in a higher inflation environment where price shocks are persistent. In this view, a shift towards lower inflation regime, brought about by more credible monetary policies, can give a rise to reduced degree of pass-through. It is worth noting that many empirical studies gave a supportive evidence to the Taylor's hypothesis, such as CHOUDHRI and HAKURA (2006), GAGNON and IHRIG (2004) and BAILLIU and FUJII (2004) to name but a few.

Another important macroeconomic determinant of pass-through is the exchange rate volatility. This latter would be positively associated with higher import price pass-through. Most of pass-through studies find that countries with low nominal exchange rate volatility have a lower ERPT. In fact the relative stability of market destination currency plays a substantial role in determining pass-through. Countries with low relative exchange rate variability would have their currencies chosen for transaction invoicing. Thereby, local currency pricing (LCP) would prevail and pass-through is less than complete. Empirically, CAMPA and GOLDBERG (2005) find that exchange rate volatility is positively associated with higher import price pass-through in 23 OECD countries, although microeconomic factors play a much more important role in determining the pass-through. For the EMU context, DEVEREUX, ENGEL, and TILLE (2003) argued that, following the formation of the EMU, the euro would become the currency of invoicing for foreign exporters (LCP). Therefore, European prices will become more insulated from exchange rate volatility and ERPT tend to be lower in such circumstance. Several Studies have tested the relevance of others macro-determinant, especially, the degree of trade openness of a country. One can expect that the more country is open, the higher is price responsiveness to exchange movements. However, results remain mitigate about the relevance of degree of openness. For instance, CHOUDHRI and HAKURA (2006) found insignificant role for the import share in their ERPT regression, while MCCARTHY (2007) provides a little evidence of a positive relationship between openness and pass-through to import price.

Table 2.1: Main ERPT studies using a Panel Cointegration analysis

STUDY	DATA	METHOD	FINDINGS
Barhouni (2006)	Annual data (1980-2003) for 24 developing countries	Measuring long run ERPT to import prices using panel data cointegration techniques. FMOLS and DOLS between-dimension estimators (Pedroni (2001)).	A higher group mean long-run ERPT coefficient: 77.2% by FMOLS, and 82.7% by DOLS. Cross-country difference in long run ERPT: by FMOLS, coefficients vary from 107% for Algeria to 42% for Chile, and by DOLS, ERPT vary from 110% for Paraguay to 43% for Singapore. Differences in ERPT are due to three macroeconomics determinants: exchange rate regimes, trade barriers and inflation regimes.
Holmes (2006)	Monthly data (1972:4- 2004:6) for 12 European Union countries	Estimation of long-run ERPT to consumer prices using DOLS between-dimension approach.	The ERPT to European Union consumer prices has declined. This decline has occurred against a background of several factors that enhanced the credibility of a low inflation regime.
de Bandt <i>et al.</i> (2008)	Disaggregated monthly data (1995-2005) for 1-digit SITC sectors for 11 euro area countries	Different panel data techniques to test for cointegration in the ERPT equation: -First generation panel cointegration tests with no cross-unit interdependence and no breaks (Pedroni (1999)) -Second generation tests with a factor structure for cross-section dependence and allowing for an individual structural break (Banerjee <i>et al.</i> (2006)).	Commodity sectors (SITC 2 and SITC 3) tend to have a higher (closer to 1) pass-through than manufacturing sectors. Strong evidence of a change in the long run ERPT behavior around the formation of the Economic and Monetary Union (EMU) or close to the period of appreciation of the euro in 2001. Long run ERPT has generally increased after these break dates especially for Italy, Portugal and Spain.
Holmes (2008)	Annual data (1971-2003) for 19 African countries.	FMOLS procedure is employed to obtain long run ERPT to import prices. Using moving window approach to test changing ERPT over time.	Long run ERPT elasticity is about 60% for the African economies. According to moving window estimates, African import prices becoming less sensitive to movements in the exchange rate over time. Decline in the long-run pass-through is accompanied by decreasing in inflation rates occurring since the mid-1990s.

In our empirical, we focus on the ERPT into import prices in the long-run, so, from econometric point of view, suitable estimation techniques must be employed. There is a crucial question about the definition of the long measure of pass-through. There are different approaches had been experimented in the empirical literature. One of the most used specifications of the long run ERPT is provided by CAMPA and GOLDBERG (2002, 2005). In these studies, the long run elasticity of pass-through is given by the sum of the coefficients on the contemporaneous exchange rate and four lags of exchange rate terms. According to DE BANDT, BANERJEE, and KOZLUK (2008), this measure is, in some extent, arbitrary, and more accurate long-run pass-through estimate must be defined. By using nonstationary panel data techniques, their study propose to restore the cointegrated long-run equilibrium in pass-through relationship (see Table 2.1 for an overview of this study). As we mentioned above, there has been an increasing use of unit root and cointegration analysis in the context of panel data. This is not surprising as panel techniques can overcome the size and power constraints associated with the use of a single time series.²

One of the most important economic theories usually tested in this context is the purchasing power parity, for which it is natural to think about long-run properties of data. However, there is a few numbers of studies has investigated the ERPT relationship within a panel data cointegration framework. In Table 2.1, we summarized the main findings of major studies in this area, namely BARHOUMI (2006), DE BANDT, BANERJEE, and KOZLUK (2007) and HOLMES (2006, 2008). Regarding country, our study is close to those of DE BANDT, BANERJEE, and KOZLUK (2007) and HOLMES (2006), which deal with some countries of the European Union. Nevertheless, our sample is larger since we consider 27 OECD countries in the first part of our analysis. Also, our country sample is more heterogeneous than the listed studies, so using PEDRONI (2001) approach is relevant since it allows the long-run cointegration relationships to be heterogeneous across countries.

²It's well-known that unit root tests have low power in small sample sizes, so adding the cross-section dimension to the time series dimension increase the power of these tests.

3. Analytical framework and Data description

3.1. Pass-Through Equation

Our approach is to use the standard specification used in the pass-through literature as a starting point (GOLDBERG and KNETTER (1997) and CAMPA and GOLDBERG (2005)). By definition, the import prices, MP_{it} , for any country i are a transformation of the export prices, XP_{it} , of that country's trading partners, using the nominal exchange rate, E_{it} (domestic currency per unit foreign currency):

$$MP_{it} = E_{it} \cdot XP_{it} \quad (2.1)$$

Using lowercase letters to reflect logarithms, we rewrite equation (2.1):

$$mp_{it} = e_{it} + xp_{it} \quad (2.2)$$

Where the export price consists of the exporters marginal cost, MC_{it} and a markup, $MKUP_{it}$:

$$XP_{it} = MC_{it} \cdot MKUP_{it} \quad (2.3)$$

In logarithms we have:

$$xp_{it} = mc_{it} + mkup_{it} \quad (2.4)$$

So we can rewrite equation (2.2) as:

$$mp_{it} = e_{it} + mc_{it} + mkup_{it} \quad (2.5)$$

Markup is assumed to have two components: (i) a specific industry component and (ii) a reaction to exchange rate movements:

$$mkup_{it} = \alpha_i + \Phi e_{it} \quad (2.6)$$

Exporter marginal costs are a function of the destination market demand conditions, y_{it} , and wages in exporting country, w_{it}^* :

$$mc_{it} = \eta_0 y_{it} + \eta_1 w_{it}^* \quad (2.7)$$

Substituting (2.6) and (2.7) into (2.5), we derive:

$$mp_{it} = \alpha_i + \underbrace{(1 + \Phi)}_{\beta} e_{it} + \eta_0 y_{it} + \eta_1 w_{it}^* \quad (2.8)$$

The structure assumes unity translation of exchange rate movements. This empirical setup permits the exchange rate pass-through, represented by $\beta = (1 + \Phi)$, to depend on the structure of competition in one industry. Exporters of a given product can decide to absorb some of the exchange rate variations instead of passing them through to the price in the importing country currency. So if $\Phi = 0$, the pass-through is complete and their markups will not respond to fluctuations of the exchange rates. This is the case when import prices are determined in the exporter's currency (producer-currency pricing or PCP). And if $\Phi = -1$, exporters decide not to vary the prices in the destination country

currency, thus they fully absorb the fluctuations in exchange rates in their own markups (LCP is prevailing).

Thus the final equation can be re-written as follows:

$$mp_{it} = \alpha_i + \beta e_{it} + \gamma y_{it} + \delta w_{it}^* + \varepsilon_{it}, \quad (2.9)$$

The most prevalent result is an intermediate case where ERPT is incomplete (but different from zero), resulting from a combination of LCP and PCP in the economy. So, there is a fraction of import prices are set in domestic currency, while the remaining prices are set in foreign currency. Thus, the extent to which exchange rate movements are passed-through to prices will depend on the predominance of LCP or PCP: the higher the LCP, the lower the ERPT, and the higher PCP, the higher ERPT.

3.2. Data description

In this study, we consider the following panel of 27 OECD countries: Australia, Austria, Belgium, Canada, Czech Republic, Denmark, Finland, France, Germany, Greece, Iceland, Ireland, Italy, Japan, Korea, Luxembourg, Netherlands, New Zealand, Norway, Poland, Portugal, Slovak Republic, Spain, Sweden, Switzerland, United Kingdom and United States. The data are quarterly and span the period 1994:1-2010:4. We use price of non-commodity imports of goods and services imports from OECD's Main Economic Outlook as a measure of the import prices, mp_{it} . From the same Data base we take the real GDP as proxy for the domestic demand, y_{it} .

To capture movements in the costs of foreign producers, W_{it}^* , that export to the domestic market, we use the same proxy adopted by BAILLIU and FUJII (2004) represented by:

$$W_{it}^* = Q_{it} \times W_{it} / E_{it} \quad (2.10)$$

Where, E_{it} , is the nominal effective exchange rate (domestic currency per unit of foreign currencies)³, W_{it} , is the domestic unit labor cost and, Q_{it} , is the real effective exchange rate. Due to data availability, we follow CAMPA and GOLDBERG (2005) by using consumer price index, P_{it} , to capture movement in production costs, assuming that prices move one-to-one to shift in wages. Taking the logarithm of each variable, we obtain the following expression:

$$w_{it}^* = q_{it} + e_{it} - p_{it} \quad (2.11)$$

Since nominal and real effective exchange rate series are trade weighted, this gives us a measure of trading-partner costs (over all partners of importing country), with each partner weighted by its importance in the importing country's trade. Data used to construct foreign producers costs - nominal effective exchange rate, Consumer prices index and real effective exchange rate - are obtained from IMF's *International Financial Statistics*.

4. Empirical methodology

4.1. Panel unit root tests for dynamic heterogeneous panels

Before testing for a cointegrating relationship, we investigate panel non-stationarity of the variables included in equation (2.9). We use the t -bar test proposed by IM et al. (2003) (henceforth IPS), which tests the null hypothesis of non stationarity.⁴ This test allows for residual serial correlation and heterogeneity of the dynamics and error variances across groups. The t -bar statistic constructed as a mean of individual ADF

³Home-currency depreciations appear as increases in the nominal effective exchange rate series.

⁴Another panel non-stationarity can be used namely HADRI (2000) test. The latter is a panel analogue of Kwiatkowski et al. (1992), tests the null hypothesis of stationarity. The Hadri test has the advantages to be suitable for panel data series with short time dimension, which is not the case of our study.

statistics and is designed to test the null that all individual units have unit roots:

$$H_0 : \rho_i = 0, \forall i$$

Against the alternative that at least one of the individual series is stationary:

$$H_1 : \begin{cases} \rho_i < 0 & \text{for } i = 1, 2, \dots, N_1 \\ \rho_i = 0 & \text{for } i = N_1 + 1, N_1 + 2, \dots, N \end{cases} \quad \text{with } 0 < N_1 \leq N$$

Where ρ_i is the coefficient of the Augmented Dickey-Fuller (ADF) regression for each individual unit⁵,

$$y_{it} = \mu_i + \rho_i y_{it-1} + \sum_{j=1}^{p_i} \varphi_{it} \Delta y_{it-j} + \gamma_i t + \varepsilon_{it}, \quad t = 1, \dots, T, \quad (2.12)$$

As we mentioned above, the IPS t -bar statistic is defined as the average of the individual ADF statistic, t_{ρ_i} , and tends to a standard normal distribution as $N, T \rightarrow \infty$ under the null hypothesis:

$$\bar{t}_{NT} = \frac{1}{N} \sum_{i=1}^N t_{\rho_i}, \quad (2.13)$$

IPS tests results are shown in Table 2.2, for both levels and first differences and with different deterministic components. In the level case, we are unable to reject the null hypothesis that all series are non-stationary in favor of the alternative hypothesis that at least one series from the panel is stationary. For tests on the first differences, we

⁵In our case all variables are assigned to y_{it} .

can see that the non-stationary null is rejected at the 5% significance level or better. We thus conclude that all variables are stationary in first difference⁶.

Table 2.2: Results for IM et al. (2003)

Variables	Level		First difference	
	Intercept	Intercept & trend	Intercept	Intercept & trend
mp_{it}	-0.5301	-1.3952	-14.0574	-17.2283
e_{it}	0.6973	-0.1440	-10.3162	-11.1541
y_{it}	-1.1128	1.0780	-7.5770	-11.0068
w_{it}^*	-1.1586	-0.9063	-10.3704	-15.7305

Note: For the IPS tests, the critical value at the 5% level is -1.81 for model with an intercept and -2.44 for model with intercept and linear time trend. Individual lag lengths are based on Akaike Information Criteria (AIC).

4.2. Tests for panel cointegration

In order to check the long run cointegrating pass-through relation, we employ PEDRONI (1999) residual-based tests. Like the IPS panel unit-root test, Pedroni’s methodology take heterogeneity into account using specific parameters which are allowed to vary across individual members of the sample.⁷ PEDRONI (1999) has developed seven tests based on the residuals from the cointegrating panel regression under the null hypothesis of non-stationarity. The first four tests (*panel v-stat*, *panel rho-stat*, *panel pp-stat*, *panel adf-stat*) are based on pooling the data along the within-dimension that are known as the *panel cointegration statistics*. The next three tests (*group rho-stat*, *group pp-stat*, *group adf-stat*) are based on pooling along the between-dimension and they are denoted *group mean cointegration statistics*. All tests are calculated using the estimated residuals from the following panel regression:

$$y_{it} = \alpha_i + \delta_{it} + \beta_1 x_{1it} + \beta_2 x_{2it} + \dots + \beta_K x_{Kit} + \varepsilon_{it}, \tag{2.14}$$

$$i = 1, \dots, N, t = 1, \dots, T, k = 1, \dots, K$$

⁶We compare the empirical statistics to the critical values given in Table 2 of IM et al. (2003) at the 5% level for $N = 25$ and $T = 70$.

⁷An alternative panel cointegration test was proposed by WESTERLUND (2007). It tests for the absence of cointegration by determining whether there exists error correction for individual panel members or for the panel as a whole. The tests are general enough to allow for a large degree of heterogeneity, both in the long-run cointegrating relationship and in the short-run dynamics, and dependence within as well as across the cross-sectional units.

In fact, both sets of test verify the null hypothesis of no cointegration:

$$H_0 : \rho_i = 1, \forall i$$

where, ρ_i is the autoregressive coefficient of estimated residuals under the alternative hypothesis ($\hat{\varepsilon}_{it} = \rho_i \hat{\varepsilon}_{it-1} + u_{it}$). We should note that the alternative hypothesis specification is different between the two sets of test:

- The *panel cointegration statistics* impose a common coefficient under the alternative hypothesis which results:

$$H_1^w : \rho_i = \rho < 1, \forall i$$

- The *group mean cointegration statistics* allow for heterogeneous coefficients under the alternative hypothesis and it results:

$$H_1^b : \rho_i < 1, \forall i$$

Pedroni has shown that the asymptotic distribution of these seven statistics can be expressed as:

$$\frac{\chi_{NT} - \mu\sqrt{N}}{\sqrt{\upsilon}} \rightarrow N(0, 1), \quad (2.15)$$

Where, χ_{NT} , is the statistic under consideration among the seven proposed, μ , and, υ , are respectively the mean and the variance tabulated in Table 2.3 of PEDRONI (1999). As shown in Table 2.3, all test statistics reject the null of no cointegration.

Table 2.3: PEDRONI (1999) Cointegration Tests Results

Tests	1994:1 - 2010:4
<i>Panel v-stat</i>	6.93854**
<i>Panel rho-stat</i>	-6.20244**
<i>Panel pp-stat</i>	-6.60297**
<i>Panel adf-stat</i>	-5.01230**
<i>Group rho-stat</i>	-5.18729**
<i>Group pp-stat</i>	-6.63478**
<i>Group adf-stat</i>	-4.72966**

Note: Except the v-stat, all test statistics have a critical value of -1.64 (if the test statistic is less than -1.64, we reject the null of no cointegration). The v-stat has a critical value of 1.64 (if the test statistic is greater than 1.64, we reject the null of no cointegration).

5. Long run ERPT estimates

Following PEDRONI (2001), we employ estimation techniques taking into account the heterogeneity of long-run coefficients. Therefore, FMOLS and DOLS Group Mean Estimator can be used to obtain panel data estimates for long run ERPT.⁸ These estimators correct the standard pooled OLS for serial correlation and endogeneity of regressors that are normally present in a long-run relationship.⁹ In our empirical analysis, we emphasis on between-dimension panel estimators. It's worth noting that the between-dimension approach allows for greater flexibility in the presence of heterogeneity across the cointegrating vectors where pass-through coefficient is allowed to vary.¹⁰ Additionally, the point estimates of the between-dimension estimator can be interpreted as the mean value of the cointegrating vectors, while this is not the case for the within-dimension estimates.¹¹ To check robustness of our result, we also reporting estimation results for the pooled OLS and fixed-effects estimators.

According to Table 2.4, long run pass-through coefficient is statistically significant with the expected positive sign, and the results are fairly robust across estimation

⁸Brief details of these methods are available in Appendix B.1.

⁹Alternatively, MARK and SUL (2003) proposed a Dynamic Ordinary Least Squares for Cointegrated Panel Data with homogeneous long-run covariance structure across cross-sectional units.

¹⁰Under the within-dimension approach pass-through elasticity would be constrained to be the same value for each country under the alternative hypothesis.

¹¹According to PEDRONI (2001) the between-group FMOLS and DOLS estimators has a much smaller size distortion than the within-group estimators.

techniques. For instance, FM-OLS estimator suggests that one percent depreciation of the nominal exchange rate increases import prices by 0.67%. As we mentioned above, pass-through equation (2.9) assume unity elasticity of import prices to exchange rate movements in order to account for complete ERPT. However, the null of unity pass-through coefficient ($H_0 : \beta = 1$) is strongly rejected through the different econometric specifications (see t -statistics reported between square brackets in Table 2.4).

Table 2.4: Panel Estimates For 27 OECD countries over 1994:1-2009:4

Variables	Dependent Variable: Import Price Index		
	Group mean FM-OLS	Group mean DOLS	Fixed effects
e_{it}	0.67*** (30.21) [16.71]	0.69*** (26.69) [16.89]	0.70*** (33.01) [10.29]
y_{it}	0.27*** (6.15)	0.20*** (6.40)	0.23*** (11.86)
w_{it}^*	0.68*** (7.09)	0.71*** (6.89)	0.214*** (8.215)

Note: Group mean FM-OLS and DOLS estimators refer to between-dimension. These estimates include common time dummies. *** indicate statistical significance at the 1 percent level. Pass-through estimates are accompanied by two t -statistics. The t -statistics in parentheses are based on the null of a zero ERPT coefficient ($H_0: \beta = 0$). The t -statistics in square brackets are based on the null of unitary elasticity ($H_0: \beta = 1$).

This is an evidence of incomplete ERPT in our sample of 27 OECD countries. On the long run, import prices do not move one-to-one following exchange rate depreciation. These results are in line with estimates in the literature of exchange rate pass-through into import prices for industrialized countries. For 23 OECD countries, CAMPA and GOLDBERG (2005) find that the average of long run ERPT is 0.64%. In this study, producer-currency pricing (or full pass-through) assumption is rejected for many countries. Using panel cointegration analysis, BARHOUMI (2006) and HOLMES (2008) reject the pass-through unity for developing countries. In accordance with the conventional wisdom that ERPT is always higher in developing than in developed countries, then a partial import prices it is expectable for OECD countries. One can think that pass-through would be complete in the long run due to the gradual full adjustment of prices (as sticky prices tend to be a short run phenomenon).¹² Nevertheless, the pricing behavior of foreign firms can prevent import prices variations following an exchange rate change. Exporters of a given product can decide to absorb some of the exchange

¹²see e.g. SMETS and WOUTERS (2002).

rate variations instead of passing them through to the price in the importing country currency. Empirically exchange rates are found to be much more volatile than prices, and then pass-through would be incomplete even in the long run. This finding is in line with the theoretical price discrimination models which assume a degree of pass-through lower than one even in the long run, as a result of PTM.

Table 2.5: Long run individual Pass-Through for 27 OECD Countries

Country	Results from FM-OLS method		
	FM-OLS	t -stat for $H_0: \beta = 0$	t -stat for $H_0: \beta = 1$
Australia	0,78*	32,7	9,04
Austria	-0,08	-0,23	3,28
Belguim	-0,04	-0,28	6,57
Canada	0,76*	18,42	5,75
Switzerland	0,39*	3,32	5,14
Czech Republic	0,95*#	10,75	0,54
Germany	0,63*	4,2	2,44
Denmark	0,28*	3,82	4,05
Spain	0,62*	4,16	2,54
Finland	-0,19	-1,49	9,53
France	0,28*	2,13	5,41
United Kingdom	0,45*	7,24	8,71
Greece	-0,11	-0,45	4,69
Ireland	0,14	1,45	8,7
Iceland	0,66*	11,44	6
Italy	0,73*#	5,25	1,92
Japan	0,44*	4,15	5,28
Korea	0,87*#	7,34	1,12
Luxembourg	0,85*#	2,44	0,43
Netherlands	0,17	1,87	9,17
Norway	0,53*	5,02	4,43
New Zealand	0,85*	16,83	2,98
Poland	0,98*#	8,01	0,14
Portugal	-0,1	-0,27	2,97
Slovak Republic	0,07	0,39	5,13
Sweden	0,48*	5,77	6,23
United States	0,38*	9,71	16,08
Mean Group panel estimation	0,67*	30,21	16,71

Note: *# implies that ERPT elasticity is significantly different from 0 (1) at the 5% level. Column (2) reports t -stat for $H_0: \beta = 0$ and column (3) reports t -stat for $H_0: \beta = 1$.

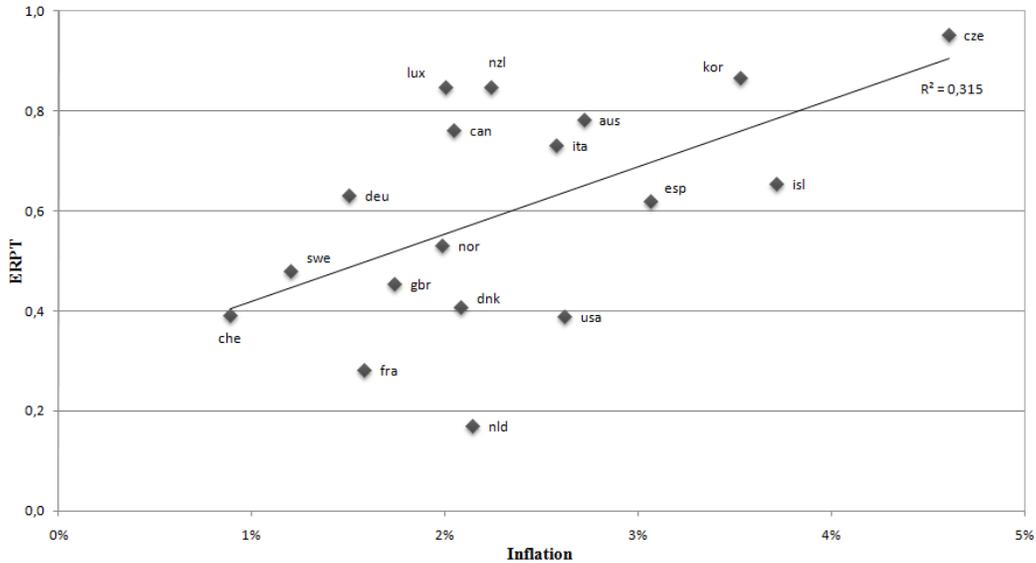
When considering individual estimates for our 27 countries, we can note a cross-country difference in the long run ERPT masked by the panel mean value. According

to Table 2.5, FM-OLS estimates show that the highest import prices reaction is in Poland by 0.98% followed by Czech Republic with 0.95%. The lowest degree of pass-through is recorded in Denmark and France with the same elasticity of 0.28% (without taking into account countries with non-significant coefficients). We can note that results are not significantly different from zero for a few numbers of countries, but it is important to mention that there is an evidence of complete pass-through for 5 out of 27 countries, namely Czech Republic, Italy, Korea, Luxembourg and Poland. This is partly corroborating CAMPA and GOLDBERG (2005) results for which producer-currency-pricing (PCP) are accepted for Poland and Czech Republic. Moreover, we can observe a low ERPT in the United States by 0.38%, which is a common result in the literature. For example, CAMPA and GOLDBERG (2005) find 41% US pass-through elasticity.

Having estimated long run ERPT coefficients, we next examine whether in line with Taylor's hypothesis there is evidence of a positive correlation between pass-through and inflation. The idea is exporters pricing strategies may be endogenous to a country's relative monetary stability. So for more stable inflation destination countries, foreign firms are willing to adopt local currency price stability (LCP) and pass-through would be incomplete. To obtain some insights on this potential positive link, we calculate year-on-year quarterly inflation rates and take the mean values over the period 1994-2009. These statistics for our 27 OECD countries are reported in Table 2.6. We should note that Japan has the lowest inflation rate with a negative value (-0.1%), while Poland experiences the highest rate exceeding 8 percent. So, in order to establish a relevant ERPT-Inflation correlation, we eliminate Japan and Poland from analysis and also countries with non-significant pass-through.

By visual inspection of Figure 2.1, we can note a positive slope arising from ERPT-Inflation. This is a strong evidence of a positive and significant association between the pass-through and the average inflation rate across countries. This finding appears overall supportive to Taylor's hypothesis. Countries with high inflation environment would experience a higher degree of pass-through. According to CAMPA and GOLDBERG (2005), although macroeconomic variables play limited role in explaining cross-country differences in ERPT, inflation rates affect significantly the extent to which exchange rate changes are "passed through" import prices.

Figure 2.1: ERPT and Inflation Correlation



Sources: Personal Calculation.

6. Macroeconomic Factors Affecting Pass-Through

Cross-country differences in the long run import prices adjustment to exchange rate would raise the question of what are the underlying determinants of pass-through. In the previous section, we have shown an important determinant of ERPT, i.e. inflation rate. Many empirical analyses have explored the influence of other macroeconomic variables such as, Exchange rate volatility and degree of openness. To pursue explanation of sources of this long run heterogeneity, we now examine some macroeconomic factors that may affect pass-through.

Three main factors are selected for this purpose: inflation rates measured as the year-on-year quarterly inflation rate; degree of openness as the percentage of import share in domestic demand (see Figure 2.2); and volatility of exchange rate changes, $\sigma_{\Delta e}$, proxied by the standard deviation of quarterly percentage changes in the nominal effective exchange rate. A summary of the average of these macroeconomic variables over 1994-2010 is given in Table 2.6. The aim of our analysis is to link those factors

to the extent of pass-through. To achieve this, we try to split our panel of countries into different groups with respect to each macroeconomic criteria, and then to estimate the ERPT for those different groups. The idea is to compare pass-through elasticity for different country regimes and to draw conclusion about the reasons of cross-country differences in ERPT into import prices.

Table 2.6: OECD Countries Statistics (1994-2009)

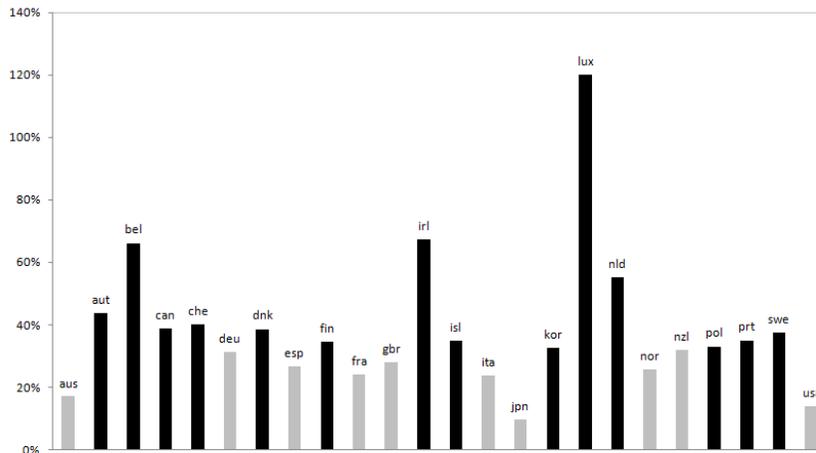
Country	Inflation Rate (%)	Openness (%)	Exchange Rate Volatility (%)
Australia	2,7	16,9	8,2
Austria	1,8	43,6	8,1
Belguim	1,8	66,1	8,6
Canada	2	38,8	4,6
Switzerland	0,9	40,2	7,6
Czech Republic	4,6	75,0	9,4
Germany	1,5	31,3	8,6
Denmark	2,1	38,6	8,2
Spain	3,1	26,4	10,7
Finland	1,4	34,5	13
France	1,6	24,1	8,2
United Kingdom	1,7	27,9	7,1
Greece	4,3	33,5	7,4
Ireland	3,7	67,4	8,3
Iceland	3,2	34,8	14,5
Italy	2,6	23,7	10,5
Japan	-0,1	9,7	8,2
Korea	3,5	32,5	13,2
Luxembourg	2	120,1	9
Netherlands	2,1	55,0	8,7
Norway	2,2	25,7	7,5
New Zealand	2	31,7	10,2
Poland	8,4	33,0	14,7
Portugal	3	35,0	9,2
Slovak Republic	6,7	77,2	9,9
Sweden	1,2	37,4	10,9
United States	2,6	13,9	5,2
Average	2,7	40,5	9,2

Note: The volatility of the exchange rate changes, $\sigma^{\Delta e}$, is computed as the standard deviation of quarterly percentage changes in the nominal effective exchange rate.

Our methodology is close to CHOUDHRI and HAKURA (2006) and BARHOUMI (2006) studies. CHOUDHRI and HAKURA (2006) classify their 71 countries into three groups based on the average of inflation rate. In their study, low, moderate and high

inflation groups are defined as consisting of countries with average inflation rates less than 10%, between 10 and 30% and more than 30%, respectively. Similarly, BARHOUMI (2006) divided a sample of 24 developing countries between high and low inflation regimes, depending on whether inflation rate is smaller or larger than 10%. However, country classification in these studies is somewhat arbitrary, in the sense that the authors used an ad hoc method to select their sample splits. In our paper, we propose to use panel threshold techniques, introduced by HANSEN (1999), to deal with the sample split problem. This methodology enables us to divide our 27 OECD countries into classes based on the value of each “macro-variables”, i.e. inflation rate, degree of openness and exchange rate volatility. To the best of our knowledge, the present paper is the only study that applying panel threshold method in this context.

Figure 2.2: Share of Imports (as a percentage of domestic demand over 1995-2009)



Source: OCDE

6.1. A single panel threshold model

HANSEN (1999) introduce a panel threshold model for a single and multiple threshold levels. Due to our small number of cross sections (27 countries), we consider the single threshold model, so that the observations can be split into two regimes depending on whether the threshold variable is above or below some threshold value. Following

HANSEN (1999), we can rewrite our pass-through equation as follow:

$$mp_{it} = \alpha_i + \beta_1 x_{it} I(q_{it} \leq \theta) + \beta_2 x_{it} I(q_{it} > \theta) + \varepsilon_{it} \quad (2.16)$$

The dependent variable of our ERPT panel threshold model is the import prices, mp_{it} , and the explanatory variables - Exchange rate, domestic demand and foreign costs - are included the vector $x_{it} = (e_{it}, y_{it}, w_{it}^*)'$. $I(\cdot)$ is an indicator function, α_i denotes the level of country i fixed-effect and ε_{it} is a zero mean, finite variance, i.i.d. disturbance. The two regimes are distinguished by different regression slopes, β_1 and β_2 , depending on whether the threshold variable q_{it} is smaller or larger than a threshold θ . If the threshold variable q_{it} is below or above a certain value, θ , then the vector of exogenous variable x_{it} has a different impact on the dependent variable, mp_{it} , with $\beta_1 \neq \beta_2$. The threshold variable q_{it} may be an element of x_{it} or a variable external to model. Effectively, in our implementation of the threshold panel method, we consider three different threshold variables - inflation rate, π_{it} , degree of openness, $open_{it}$, and exchange rate volatility, $\sigma_{it}^{\Delta e}$ - which are not belonging to explanatory variables of the pass-through equation. Thus, we will estimate equation (2.16) for our different threshold variables, $q_{it} = \pi_{it}, open_{it}, \sigma_{it}^{\Delta e}$.

The determination of the estimated threshold, $\hat{\theta}$, is based on two steps procedure using ordinary least squares (OLS) method¹³. In the first step, for any given threshold, θ , the sum of square errors is computed separately. In the second step, by minimizing of the sum of squares of errors, $S_1(\theta)$, the estimated threshold value, $\hat{\theta}$ is obtained and the residual variance, $\hat{\sigma}^2$, is saved. To check whether the threshold is in fact statistically significant, the null hypothesis of no threshold effect is tested: $H_0 : \beta_1 = \beta_2$. The likelihood ratio test of H_0 is based on the following F -statistics: $F_1 = (S_0 - S_1(\hat{\theta})) / \hat{\sigma}^2$, where S_0 and $S_1(\hat{\theta})$ are sum of squared errors under null and alternative hypotheses, respectively. The asymptotic distribution of F_1 is non-standard. HANSEN (1999) propose to use a bootstrap procedure to compute the p -value for F_1 under H_0 .

¹³Estimation techniques for panel threshold model is given in the Appendix B.1.3 with more details.

Once a significant single threshold is found, we can estimate the pass-through coefficient for each regime. For the purpose of our analysis, we use the estimated threshold to divide our country sample into different groups with respect to their macroeconomic environment (inflation level, degree of openness and exchange rate volatility)¹⁴. Then, we estimate the ERPT elasticity for each class of countries in order to make a comparison between different regimes.

6.2. Estimation of a single threshold

The estimation results of the threshold levels for each of our macro-determinant are reported in Table 2.7. Also, we give the plots of sum of squared residuals for the different threshold variables (see Figure 2.3). When we consider inflation rate as threshold variables, ($q_{it} = \pi_{it}$), we find a threshold level close to 2% ($\theta_{\pi} = 0,019$). The test for a single threshold is significant with a bootstrap p -value of 0,04. Given this threshold value, we can define two groups of countries based on *inflation-regime*, i.e. with respect to the average of inflation rate. Thus, we consider countries with mean of inflation equal or less than 2% as *low inflation* countries, while countries with inflation mean more than 2% as *moderate inflation* countries¹⁵. According to this classification, we obtain 12 low inflation countries and 15 countries with moderate inflation-regime (see Table 2.8).

The next threshold variable considered in pass-through equation is the degree of openness (Figure 2.2 gives the import shares of our 27 OECD countries). According to Table 2.7, the estimated threshold value is 31,8% of import share, but the presence of a single threshold is insignificant according to bootstrapped p -value (0,26). Nevertheless, this threshold value is still the best point to consider for splitting our sample with the respect to the degree of openness (see figure 2.3). Thus, we will consider countries characterized by degree of openness less than 32% as *less open* countries, while countries having import share larger than 32% will be defined as *more open* countries. This gives us 10 less open countries and 17 countries with degree of openness more than 32%.

¹⁴We follow the same strategy of HANSEN (1999) who used the threshold values to split his sample of 565 US firms into *low debt* and *high debt* firms.

¹⁵The term of *moderate inflation* is used instead of *high inflation* since we don't have double-digit inflation countries in our sample of 27 OECD countries.

Table 2.7: HANSEN (1999) test for a single threshold
$$mp_{it} = \alpha_i + \beta_1 x_{it} I(q_{it} \leq \theta) + \beta_2 x_{it} I(q_{it} > \theta) + \varepsilon_{it}$$

Inflation rate: ($q_{it} = \pi_{it}$)	
Threshold value ($\hat{\theta}_\pi$)	0.019
F-test	168.03
Bootstrapped p-values	(0.040)
Degree of openness: ($q_{it} = open_{it}$)	
Threshold value ($\hat{\theta}_{open}$)	0.318
F-test	63.902
Bootstrapped p-values	(0.260)
Exchange rate volatility: ($q_{it} = \sigma_{it}^{\Delta e}$)	
Threshold value ($\hat{\theta}_{\sigma^{\Delta e}}$)	0.082
F-test	78.738
Bootstrapped p-values	(0.010)

Note: Table reports threshold estimates ($\hat{\theta}$), F-test of the null hypothesis of no threshold effect and bootstrapped p-values obtained from 1000 bootstrap replications.

Finally, the last criterion which can explain differences in pass-through elasticity is the exchange rate volatility. Different sort of proxies are used in the ERPT literature. For instance, CAMPA and GOLDBERG (2005) take the average of the monthly squared changes in the nominal exchange rate. For MCCARTHY (2007) exchange rate volatility is measured by the variance of the residuals from the exchange rate equation in the VAR. In our empirical analysis, we adopt the same exchange rate volatility proxy employed by BARHOUMI (2006) and compute exchange rate volatility as the standard deviation of quarterly percentage changes in the exchange rate, $\sigma^{\Delta e}$.¹⁶ According to Hansen's single threshold test, we find a significant threshold value equal to 0.082 (see Table 2.7). Accordingly, we will call countries for whom the mean of exchange rate volatility is less than 8.2% as *less volatility* countries, and the sub-sample of countries having $\sigma^{\Delta e}$ more than 8.2% as *high volatility* countries. We count 11 low volatility countries and 16 high volatility countries (see Table 2.8).

¹⁶To obtain exchange rate volatility series, we start by computing the standard deviation of changes in exchange rate for the first quarter 1994:1 during the last five years and, then, we slid forward this window quarter by quarter throughout our estimation period (1994-2010).

Figure 2.3: Threshold levels according to sum of squared residuals

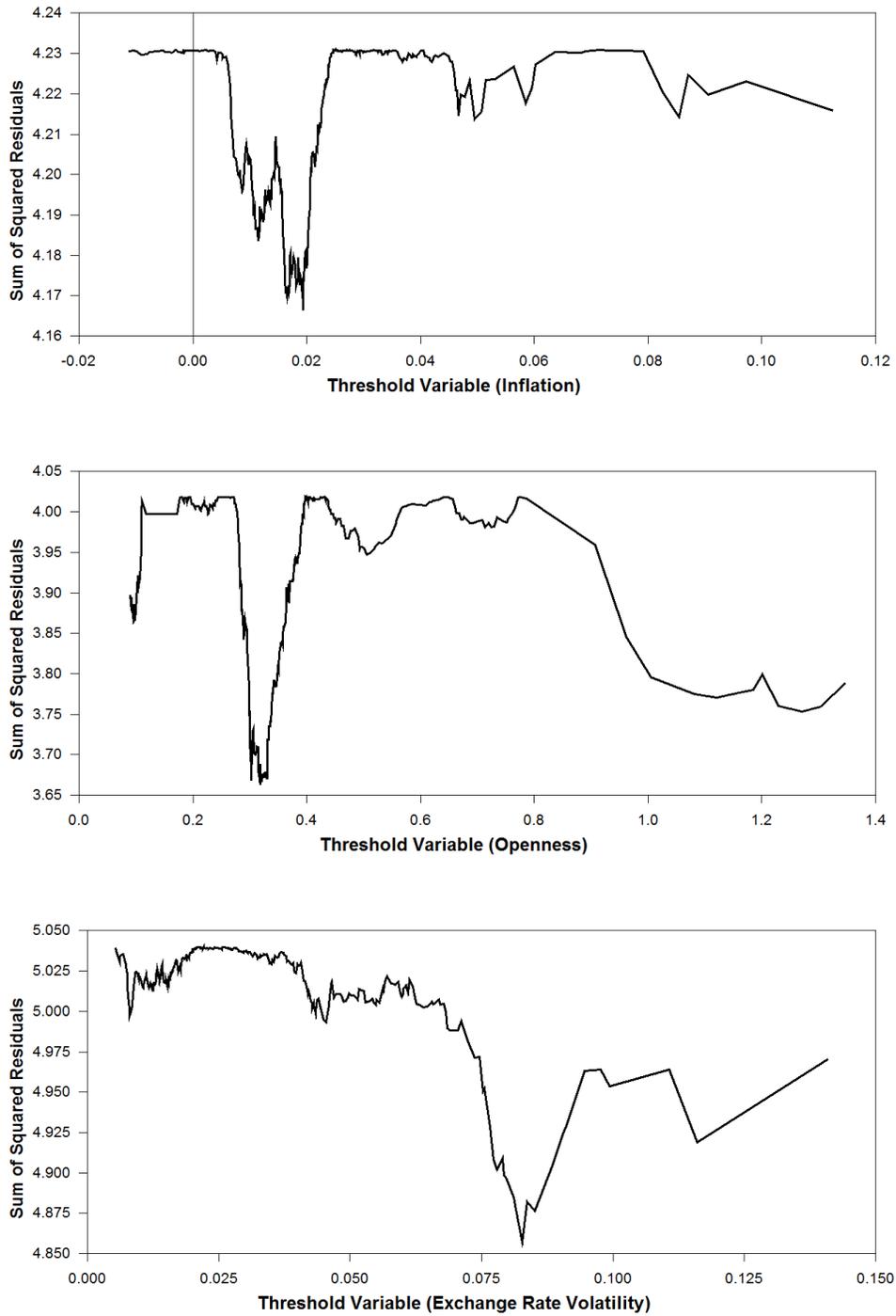


Table 2.8: Country Classification

Inflation Regime		Degree of Openness			Exchange Rate Volatility			
Low Inflation	Moderate inflation	Less Open	More Open	Low Volatility	High Volatility			
Austria	Australia	Portugal	Australia	Austria	Netherlands	Australia	Belguim	Poland
Belguim	Czech Republic	Slovak Republic	Germany	Belguim	Poland	Austria	Czech Republic	Portugal
Canda	Denmark	United States	Spain	Canda	Portugal	Canda	Germany	Slovak Republic
Switzerland	Spain		France	Switzerland	Slovak Republic	Switzerland	Spain	Sweden
Germany	Greece		United Kingdom	Czech Republic	Sweden	Denmark	Finland	
Finland	Ireland		Italy	Denmark		France	Ireland	
France	Iceland		Japan	Finland		United Kingdom	Iceland	
United Kingdom	Italy		Norway	Greece		Greece	Italy	
Japan	Korea		New Zealand	Ireland		Japan	Korea	
Luxembourg	Netherlands		United States	Iceland		Norway	Luxembourg	
New Zealand	Norway			Korea		United States	Netherlands	
Sweden	Poland			Luxembourg			New Zealand	
12 countries	15 countries		10 countries	17 countries		11 countries	16 countries	

Note: Last line denotes number of countries in each class. The volatility of the exchange rate changes, $\sigma^{\Delta e}$, is computed as the standard deviation of quarterly percentage changes in the exchange rate.

6.3. Regime dependence of ERPT

Following countries classification, now we must perform estimation for each panel group of countries. So before applying FM-OLS and DOLS estimators, we proceed by testing panel unit root for individual series within each group (high and low inflation, more and less open countries, and more and less exchange rate volatility). Results from IPS tests (reported in Appendix B.2.1) show that most of variables are $I(1)$. Then, we provide the presence of cointegration relationship by using Pedroni cointegration tests for different sub-sample panel of countries (Appendix B.2.2). Almost all of tests lead us to reject the null of non-cointegration.

Estimates of long-run ERPT for each group of countries reported in Table 2.9. We begin with the inflation rate as a macro-determinant of the extent of pass-through. In view of results, low inflation countries experience long run import prices elasticity equal to 0.53% by FM-OLS. While one percent exchange rate depreciation causes an increase in import prices by 0.75% in high inflation countries. Results remain robust when using DOLS method. Thus, ERPT is found to be higher in high inflation environment countries. It is evident that this finding corroborates the convention wisdom of the positive link between Inflation and pass-through (TAYLOR (2000)). That is, countries with higher rates of inflation should have higher rates of pass-through of exchange rates into import prices. Our results provide an evidence of regime-dependence of ERPT with respect to inflation environment and this latter would be an important source of heterogeneity in pass-through across countries.

For our second macro-determinant, i.e. import share, one can expect a positive connection between openness and pass-through: the more a country is open, the more import prices respond to exchange rate fluctuations. According to our results this positive link seems to be weak. Both FM-OLS and DOLS show a long-run ERPT of roughly 0.56% in less open economies, which is little smaller than in the more open ones (0.68% by FM-OLS). The 95% confidence band shows that the extent of pass-through seems to do not differ strongly between the two groups of country, especially according to DOLS estimators. As we mentioned above, there is no conclusive empirical results in the literature about the relevance of degree of openness. For nine developed countries, MCCARTHY (2007) shows that association is not significant between import

share and pass-through¹⁷. However, BARHOUMI (2006) found a positive correlation of pass-through-openness in panel cointegration framework. The main difference with our analysis is that the measure of openness used in BARHOUMI (2006) is the tariffs barriers. The author found that lower tariff barriers countries experience a higher long run pass-through than higher tariff barriers.

Table 2.9: Long run Pass-Through Estimates for different country regime

	Inflation Regime		Degree of openness		Exchange rate volatility	
	Low Inflation	High Inflation	Less Open	More Open	Less volatile	More volatile
FMOLS	0,53** [0,49 0,57]	0,75** [0,70 0,81]	0,57** [0,53 0,60]	0,68** [0,62 0,57]	0,47** [0,43 0,52]	0,79** [0,74 0,84]
DOLS	0,51** [0,46 0,55]	0,82** [0,76 0,89]	0,56** [0,52 0,60]	0,66** [0,58 0,75]	0,39** [0,35 0,43]	0,74** [0,69 0,79]
	12 countries	15 countries	10 countries	17 countries	11 countries	16 countries

Note: ** indicate statistical significance at the 5 percent level. 95% confidence intervals are reported between square brackets.

It is worthwhile to note that ROMER (1993) provide an indirect channel, whereby openness is negatively correlated with inflation of consumer prices. In this study, he explains that real depreciation caused by unexpected expansionary monetary policy might be harmful in more open economies (with absence of binding precommitment), thus the benefits of expansion is negatively correlated with the degree of openness. Therefore, monetary authorities in more open countries would expand less which results in lower inflation rates. Nevertheless, our empirical analysis is concerned with pass-through to import prices and not to consumer prices. The main explanation of the negative effect of openness on import prices ERPT is that the greater openness of a country is an indicative of increased competitive pressures between foreign and local producers. In this case, foreign firms are willing to accept adjustments to their markup in order to maintain market share, and the extent of pass-through to import prices would be lower.

Finally, we raise the question about the relevance of exchange rate volatility in explaining the long run pass-through. In fact, it is expected that import prices responsiveness would be higher when volatility of exchange rate is larger. As pointed

¹⁷Similarly, CHOUDHRI and HAKURA (2006) found a little evidence of a positive relationship between ERPT to consumer prices and openness

by DEVEREUX and ENGEL (2002), the relative stability of market destination currency plays a substantial role in determining pass-through. Countries with low relative exchange rate variability would have their currencies chosen for transaction invoicing. Thereby, local-currency pricing (LCP) would prevail and pass-through is less than complete. In view of our results, pass-through elasticity is about 0.40% in less volatility exchange rate countries, but import prices increase by 0.74% following one percent nominal depreciation in high volatility countries (according to DOLS estimates). There is significant difference between the two groups, and results are robust across FM-OLS and DOLS estimates. Empirically, this finding is consistent with CAMPA and GOLDBERG (2005) for whom higher home currency volatility is significantly associated with lower ERPT.

It is important to mention that this positive link between is not as obvious as one would think. In his VAR Study, MCCARTHY (2007) suggest that that pass-through should be less in countries where the exchange rate has been more volatile. The author argued that greater home currency volatility may make exporters more willing to adjust profit margins, which reduces measured pass-through. In his panel of developing countries, BARHOUMI (2006) gives support to this intuition. He obtains a lower pass-through for fixed exchange rate regime countries which are defined as panel group with less volatile exchange rate.

7. Has ERPT declined in the Euro Area?

In the final part of the paper, it is useful to focus on the case of the euro area (EA) members by taking the following sub-sample of 12 countries: Austria, Belgium, Germany, Spain, Finland, France, Greece, Ireland, Italy, Luxembourg, Netherlands and Portugal. It is important to emphasize that the formation of the euro area is likely to have an important impact on ERPT. This could be true since the launch of the monetary union in January 1999 is seen as a shift in both monetary policy regime and competition conditions.

Empirically, little is said about the EA long run ERPT to import prices in context of panel cointegration analysis. As summarized in Table 2.1, HOLMES (2006) examine

the pass-through question for 12 European Union members, and his sample involves countries not belonging to the monetary union such as Denmark, UK and Sweden. Also his analysis is concerned with ERPT to consumer prices, and not with the first stage of pass-through, i.e. ERPT to import prices. For 11 euro area countries, DE BANDT, BANERJEE, and KOZLUK (2007) deal with the micro level of pass-through rather than focusing on aggregate prices reactions, by considering the 1-digit SITC import prices sectors. Therefore, the aim of this section is slightly different from these two studies. We investigate the degree to which exchange rate variations are transmitted into import prices on the long run and at the aggregate level for 12 countries of the EMU.

It is commonly agreed that the observed decline in pass-through has coincided with general reduced and stable inflation rates in many countries. And consequently, more credible monetary policy regime is seen as a main determinant factor that insulating prices volatility to home currency depreciation. Since the end of Bretton-woods era, European countries have experienced various macroeconomic developments notably in terms of monetary policy and exchange rate regime. This was started with the *snake in the tunnel* period, followed by the entering to the ERM within the EMS, which has led later to the launch of the EMU and the adoption of the euro in January 1999. During this long period of time, it is naturally to see that European countries have gone through different inflation regimes as reported in Table (2.10).

There has been a steady decline in the mean inflation in our sample of 12 EA countries, which has fallen from 11.4% during the European Snake period to 2.4% over the last decade. It is expectable that this inflation behavior has influenced the extent to which prices respond to exchange rate changes. With referring to TAYLOR (2000), ERPT tend to decline in countries where monetary policy shifted strongly towards stabilizing inflation. Thereby, we try to address this question, and investigate whether or not ERPT has been diminished since the demise of Bretton-Woods.

Table 2.10: Inflation Rates in the EMU (from 1972:2 to 2010:4)

Country	European Snake 1972:2-1979:1	EMS 1979:2-1990:2	1 st & 2 nd stage of EMU 1990:3-1998:4	3 rd stage of EMU 1999:1-2010:4
Austria	6,8	3,8	2,6	1,8
Belguim	8,3	4,8	2,2	2
Germany	5,1	3	2,8	1,5
Spain	16,6	10,4	4,6	3,3
Finland	12,3	7,2	2	1,6
France	9,7	7,4	2	1,6
Greece	14,5	19,5	11,7	3,1
Ireland	13,9	9,4	2,3	2,9
Italy	14,4	11,3	4,3	2,2
Luxembourg	7,3	4,7	2,2	2,2
Netherlands	7,7	2,9	2,5	2,1
Portugal	20,6	17,9	6	2,5
Average	11,4	8,5	3,8	2,2

Source: OCDE

Now we take a long time series quarterly data from 1973:2 to 2010:1, and then proceed by splitting this sample period into four sub-periods defined according to different exchange rate and monetary policy arrangements: First period corresponds to the snake in the tunnel phase from 1972:2 to 1979:1; the second is the SME period from 1979:2 to 1990:2; third period record the launch of the first stage of the EMU and involves also the second stage which finish in 1998:4 ; and the last period corresponds to the formation of the euro area in 1999:1 and lasts until 2010:4. This empirical analysis consist of estimating equation (2.9) for each of these four sub-periods, in order to assess the ERPT development through different regimes where inflation rate have been considerably declined.

For each sub-period, we conduct IPS tests to check the presence of panel unit root in variables series. As seen in Table 2.11, we are unable to reject the null of non-stationarity for most of series in level. Also, we test for cointegration relationships between the variables for the four sub-periods. *Group PP* and *Group ADF* Pedroni tests reject the null of non-cointegration in favor of the alternative of cointegration for all countries.

Table 2.11: IPS and Pedroni tests for EA sub-sample

	European Snake 1972:2-1979:1	EMS 1979:2-1990:2	1 st & 2 nd stage of EMU 1990:3-1998:4	3 rd stage of EMU 1999:1-2010:4
IPS Unit Root tests				
mp_{it}	-1.5022	0.4160	-0.3307	0.0195
e_{it}	0.2934	0.4810	-1.0624	-0.2201
y_{it}	-1.2372	1.5465	-0.7894	0.3976
w_{it}^*	-0.5819	-0.5473	0.0887	-1.0461
Δmp_{it}	-2.450***	-2.579***	-4.1707	-4.8766
Δe_{it}	-2.657***	-2.558	-2.4721	-3.2593
Δy_{it}	-3.310***	-3.1612	-2.7697	-3.0879
Δw_{it}^*	-3.398***	-4.690***	-4.6786	-6.1484
Pedroni Cointegration tests				
<i>Group pp-stat</i>	-3.268**	-4.203**	-6.093**	-2.271**
<i>Group adf-stat</i>	-3.141**	-4.779**	-4.225**	-3.110**

Note: For the IPS tests, the critical value at the 5% level is -1.81 for model with an intercept and -2.44 for model with intercept and linear time trend. Individual lag lengths are based on Akaike Information Criteria (AIC). For Pedroni tests, Group pp and Group ADF test statistics have a critical value of -1.64. If the test statistic is less than -1.64, we reject the null of no cointegration.

Now we return to the estimation results reported in Table 2.12. As expected, both FM-OLS and DOLS estimators give a strong evidence of a decline of long run ERPT throughout sub-sample periods. During the *snake in the tunnel* period, import prices responsiveness was higher equaling 0.90% following one percent currency depreciation (by FM-OLS), and this is the utmost pass-through elasticity recorded among the four sub-periods. It is interesting to note that this highest ERPT coefficient occurs in period where mean of inflation rates exceeding the 10% percent in our sample. The moving to SME arrangement does not change considerable the degree to which exchange rate movements affect import prices. Over the period 1979:2-1990:2, ERPT still a little bit higher and upper to 0.80% referring to DOLS estimate. Inflation rates remain higher during this sub-period with more than 8% in average. For the third period (first and second EMU stage), exchange rate depreciation is transmitted in a lesser extent, that is, import prices increase by only 0.60%. It is worthwhile noting that this lowering in pass-through coincides with a substantial fall in the mean of EA inflation rate (3.8%). This result advocates for plausible association between inflation and ERPT. Similarly, since the adoption of the common currency in 1999 pass-through remains lower than European snake and SME periods. We obtain 0.50% (0.53%) by FM-OLS (by DOLS)

as import prices reactions. We note that pass-through elasticities are not quite different in comparison with 1990:3-1998:4 sub-period (Figure 2.4).

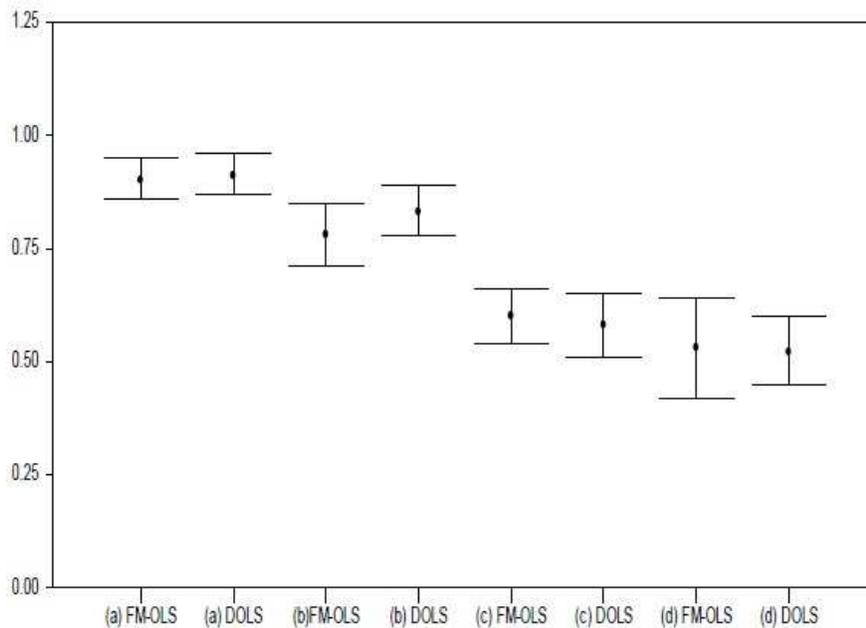
Table 2.12: Long run ERPT into imports prices in the EMU

	European Snake 1972:2-1979:1	EMS 1979:2-1990:2	1 st & 2 nd stage of EMU 1990:3-1998:4	3 rd stage EMU 1999:1-2010:4
FM-OLS	0,90** [0,86 0,95]	0,78** [0,71 0,85]	0.60** [0,54 0,66]	0.53** [0,42 0,64]
DOLS	0,91** [0,87 0,96]	0,83** [0,78 0,89]	0.58** [0,51 0,65]	0.52** [0,45 0,60]
Inflation	11,4	8,5	3,8	2,2

Note: ** indicate statistical significance at the 5 percent level. 95% confidence intervals are reported between square brackets.

Given these results, we notice that the broad decline in long run ERPT is concordant to the steady decline in the mean inflation in our 12 euro area countries. In the light of Taylor hypothesis, it is arguable that this behavior in inflation rate has gave rise a decline in the degree of pass-through. Consequently, a possible positive link between ERPT and inflation can be established in our sub-sample of euro area.

These findings are so convincing since the two macro-determinants, i.e. inflation rates and exchange rate volatility have become more stable throughout the whole sample period (1972:2 to 2010:1). Except few troubling events, European currencies have achieving more stability among each other along the different monetary policy transition which ended with the formation of the euro area in 1999. In the same way, EMU members have gained more credibility through a sustained commitment to maintaining low inflation, and this has been enforced by the explicit primary objective of European Central Bank (ECB), i.e. the price stability. Our findings are in line with the suggestion of DEVEREUX, ENGEL, and TILLE (2003) who argue that the euro would become the currency of invoicing (LCP). As a result European imports prices would become more insulated from exchange rate movements, and thereby ERPT would fall upon the introduction of the euro.

Figure 2.4: Long run ERPT Estimates for the Euro Area

Note : (a) European snake period 1972:2-1979:1 ;
 (b) SME period 1979:2-1990: 2 ;
 (c) 1st and 2nd stage of EMU 1990:3-1998:4 ;
 (d) 3rd stage EMU 1999:1-2010:1.

8. Conclusion

This paper has examined the long run exchange rate pass-through (ERPT) into import prices using panel cointegration approach. We first provide a strong evidence of incomplete ERPT in sample of 27 OECD countries. On the long run, import prices do not move one to one following exchange rate depreciation. Both FM-OLS and DOLS estimators show that pass-through elasticity does not exceed 0.70%. These results are in line with estimates in the literature of exchange rate pass-through into import prices for industrialized countries. When considering individual estimates for our panel of 27 countries, we can note a cross-country difference in the long run ERPT, with the highest import prices reactions are recorded in Poland by 0.98% followed by Czech Republic with 0.95%. It is important to mention that there is an evidence of complete pass-

through for 5 out of 27 countries, namely Czech Republic, Italy, Korea, Luxembourg and Poland. The cross-country differences in the pass-through lead us to the question of what are the underlying determinants of pass-through. According to the individual coefficients, there is a significant positive correlation between inflation rates and the extent to which exchange rate variations are passed through import prices.

Then, when split our sample in two inflation country regime, we find that high inflation countries have experienced a higher degree of ERPT than lower inflation ones. These findings are in line with Taylor's hypothesis. Another potential source of cross-country differences is home currency volatility. In view of our results, import prices responsiveness would be lower in countries with less volatile exchange rate. This can be explained by foreign firms' behaviors which are willing to set their prices in stable currency country (local currency pricing (LCP)). We can mention that we find a weak evidence of a positive link between degree of openness and ERPT which is commonly agreed in the pass-through literature.

In the final part of our analysis, we focus on the case of the European Monetary Union (EMU) by taking a sub-sample of 12 euro countries. Our goal is to assess the behavior of ERPT since the collapse of Breton-Woods era and to try to relate it to the change in the inflation environment. As a result, we find a steady decline in the degree of pass-through throughout the different exchange rate arrangements: ERPT elasticity was close to unity during the *snake-in-the tunnel* period while it is about 0.50% since the formation of the euro area. Finally, it is important to emphasize that the observed decline in pass-through to import prices was synchronous to the shift towards reduced inflation regime in our sample of countries. There is a broad downward tendency for both inflation and ERPT. This can give a further evidence of the positive correlation between price stability regime and the extent of pass-through.

Appendix B

B.1. Estimation methods

B.1.1. FM-OLS Mean Group Panel Estimator: PEDRONI (2001)

We consider the following fixed effect panel cointegrated system:

$$y_{it} = \alpha_i + x'_{it}\beta + \varepsilon_{it}, \quad t = 1, \dots, T, \quad (\text{B.1})$$

x'_{it} , can in general be a m dimensional vector of regressors which are integrated of order one, that is:

$$x_{it} = +x_{it-1} + u_{it}, \forall i \quad (\text{B.2})$$

Where the vector error process $\xi_{it} = (\varepsilon_{it}, u_{it})'$ is stationary with asymptotic covariance matrix:

$$\Omega_{it} = \lim_{T \rightarrow \infty} E \left[T^{-1} \left(\sum_{t=1}^T \xi_{it} \right) \left(\sum_{t=1}^T \xi'_{it} \right) \right] = \Omega_i^0 + \Gamma_i + \Gamma_i' \quad (\text{B.3})$$

Ω_i^0 , is the contemporaneous covariance and, Γ_i , is a weighted sum of autocovariances.

The long run covariance matrix is constructed as follow: $\begin{bmatrix} \Omega_{11i} & \Omega'_{21i} \\ \Omega_{21i} & \Omega_{22i} \end{bmatrix}$, where, Ω_{11i} , is the scalar long run variance of the residual, ε_{it} , and, Ω_{22i} , is the long run covariance among the, u_{it} , and, Ω_{21i} , is vector that gives the long run covariance between the residual, ε_{it} , and each of the u_{it} .

For simplicity, we will refer to, x_{it} , as univariate. So according to PEDRONI (2001), the expression for the group-mean panel FM-OLS estimator (for the between dimension) is given as:

$$\hat{\beta}_{GFM} = N^{-1} \sum_{i=1}^N \left(\sum_{t=1}^T (x_{it} - \bar{x}_i)^2 \right)^{-1} \times \left(\sum_{t=1}^T (x_{it} - \bar{x}_i) y_{it}^* - T \hat{\gamma}_i \right) \quad (\text{B.4})$$

Where $y_{it}^* = (y_{it} - \bar{y}_i) - \frac{\hat{\Omega}_{21i}}{\hat{\Omega}_{22i}} \Delta x_{it}$, and $\hat{\gamma}_i \equiv \hat{\Gamma}_{21i} - \Omega_{21i}^0 - \frac{\hat{\Omega}_{21i}}{\hat{\Omega}_{22i}} (\hat{\Gamma}_{22i} - \Omega_{22i}^0)$, with $y_i = \frac{1}{T} \sum_{t=1}^T y_{it}$ and $x_i = \frac{1}{T} \sum_{t=1}^T x_{it}$ refer to the individual specific means.

The Pedroni between FM-OLS estimator, $\hat{\beta}_{GFM}$, is the average of the FMOLS estimator computed for each individual, $\hat{\beta}_{FM,i}$, that is:

$$\hat{\beta}_{GFM} = N^{-1} \sum_{i=1}^N \hat{\beta}_{FM,i} \quad (\text{B.5})$$

The associated t -statistic for the between-dimension estimator can be constructed as the average of the t -statistic computed for each individuals of the panel:

$$t_{\hat{\beta}_{GFM}} = N^{-1/2} \sum_{i=1}^N t_{\hat{\beta}_{FM,i}} \quad (\text{B.6})$$

Where $t_{\hat{\beta}_{FM,i}} = \left(\hat{\beta}_{FM,i} - \beta_0 \right) \left(\hat{\Omega}_{11i}^{-1} \sum_{t=1}^T (x_{it} - \bar{x}_i)^2 \right)^{1/2}$.

B.1.2. DOLS Mean Group Panel Estimator: PEDRONI (2001)

The DOLS regression can be employed by augmenting the cointegrating regression with lead and lagged differences of the regressors to control for endogenous feedback effects. Thus, we can obtain from the following regression:

$$y_{it} = \alpha_i + \beta_i x_{it} + \sum_{k=-K_i}^{K_i} \gamma_{it} \Delta x_{it-k} + \varepsilon_{it}, \quad (\text{B.7})$$

The group-mean panel DOLS estimator is construct as:

$$\hat{\beta}_{GD} = N^{-1} \sum_{i=1}^N \left(\sum_{t=1}^T Z_{it} Z'_{it} \right)^{-1} \left(\sum_{t=1}^T Z_{it} \tilde{y}_i \right) \quad (\text{B.8})$$

Where $Z_{it} = (x_{it} - \bar{x}_i, \Delta x_{it-K}, \dots, \Delta x_{it-K})$ is a the $2(K+1) \times 1$ vector of regressors and $\tilde{y}_i = y_{it} - \bar{y}_i$.

The DOLS estimator for the i^{th} member of the panel is written as:

$$\hat{\beta}_{D,i} = \left(\sum_{t=1}^T Z_{it} Z'_{it} \right)^{-1} \left(\sum_{t=1}^T Z_{it} \tilde{y}_i \right) \quad (\text{B.9})$$

So that the between-dimension estimator can be constructed as

$$\hat{\beta}_{GD} = N^{-1} \sum_{i=1}^N \hat{\beta}_{D,i} \quad (\text{B.10})$$

If the long-run variance of the residuals from the DOLS regression (B.7) is:

$$\sigma_i^2 = \lim_{T \rightarrow \infty} E \left[T^{-1} \left(\sum_{t=1}^T \varepsilon_{it} \right)^2 \right] \quad (\text{B.11})$$

According to Pedroni, the associated t-statistic for the between-dimension estimator can be constructed as:

$$t_{\hat{\beta}_{GD}} = N^{-1/2} \sum_{i=1}^N t_{\hat{\beta}_{D,i}} \quad (\text{B.12})$$

$$\text{Where } t_{\hat{\beta}_{D,i}} = \left(\hat{\beta}_{D,i} - \beta_0 \right) \left(\hat{\sigma}_i^{-2} \sum_{t=1}^T (x_{it} - \bar{x}_i)^2 \right)^{1/2}.$$

B.1.3. Estimation of Panel Single Threshold Model: HANSEN (1999)

Equation (2.16) in the text can be written as follows:

$$y_{it} = \alpha_i + \beta' x_{it}(\theta) + \varepsilon_{it}, \quad (\text{B.13})$$

Where y_{it} is the dependent variable, $x_{it}(\theta) = \begin{pmatrix} x_{it}I(q_{it} \leq \theta) \\ x_{it}I(q_{it} > \theta) \end{pmatrix}$ is a k -dimensional vector of exogenous variables and $\beta = (\beta_1', \beta_2')$.

After removing the individual-specific means, α_i , using the within transformation estimation techniques, the OLS estimator of β is given by:

$$\hat{\beta}(\theta) = (X^*(\theta)' X^*(\theta))^{-1} X^*(\theta)' Y^* \quad (\text{B.14})$$

Where X^* and Y^* denote the stacked data over all individuals after removing the individual specific means.

The vector of regression residuals is $\hat{\varepsilon}^*(\theta) = Y^* - X^*(\theta)\hat{\beta}(\theta)$ and the sum of squared errors can be written as:

$$S_1(\theta) = \hat{\varepsilon}^*(\theta)' \hat{\varepsilon}^*(\theta) = Y^{*'} (I - X^*(\theta)'(X^*(\theta)'X^*(\theta))^{-1}X^*(\theta)')Y^* \quad (\text{B.15})$$

In a second step HANSEN (1999) recommend the estimation of the threshold θ by least squares which is achieved by minimization of the concentrated sum of squared errors $S_1(\theta)$. Then, the least squares estimators of $\hat{\theta}$ is given by:

$$\hat{\theta} = \underset{\theta}{\operatorname{argmin}} S_1(\theta) \quad (\text{B.16})$$

Hence, the resulting estimate for the slope coefficient is obtained by $\hat{\beta} = \hat{\beta}(\hat{\theta})$. The residual vector is $\hat{\varepsilon}^* = \hat{\varepsilon}^*(\hat{\theta})$ and residual variance is defined as:

$$\hat{\sigma}^2 = \frac{1}{N(T-1)} \hat{\varepsilon}^{*'} \hat{\varepsilon}^* = \frac{1}{N(T-1)} S_1(\hat{\theta}) \quad (\text{B.17})$$

B.2. Stationarity and cointegration tests for different regimes

B.2.1. Panel unit root tests

Table B.1: IPS tests for different country regime

	Level		First difference	
	Intercept	Intercept & trend	Intercept	Intercept & trend
Low Inflation				
mp_{it}	-0.1405	-1.0594	-7.9740	-9.1933
e_{it}	-0.8485	-0.8553	-10.9606	-8.0266
y_{it}	-0.1405	3.5652	-9.3488	-8.7950
w_{it}^*	0.3445	-0.2818	-8.8177	-9.5902
High Inflation				
mp_{it}	-0.6878	-0.2549	-5.9312	-8.2145
e_{it}	2.2388	0.0321	-7.6749	-7.8984
y_{it}	-0.5381	3.1936	-6.7155	-6.6649
w_{it}^*	0.2137	-1.1218	-5.3517	-6.7586
Low Openness				
mp_{it}	-0.6883	-1.0806	-4.9044	-4.9137
e_{it}	0.6127	-1.6987	-7.0398	-9.5322
y_{it}	-1.4890	-0.5947	-3.8447	-3.7240
w_{it}^*	2.2661	-0.8751	-3.9590	-3.9204
High Openness				
mp_{it}	0.0553	-0.2854	-3.7535	-3.7441
e_{it}	2.6988	0.6116	-6.2117	-6.7251
y_{it}	0.1784	1.2138	-6.2556	-5.7448
w_{it}^*	0.5523	-0.3322	-6.5015	-3.2179
Low Volatility				
mp_{it}	0.1393	-0.5981	-5.1119	-5.7440
e_{it}	1.4496	0.3933	-5.3775	-4.3244
y_{it}	-1.7040	4.0617	-0.4306	-8.9381
w_{it}^*	1.4477	-0.9389	-6.5228	-3.9717
High Volatility				
mp_{it}	-1.0527	-0.2813	-3.5419	-5.6684
e_{it}	0.4843	-0.7306	-2.9523	-7.5928
y_{it}	0.8293	2.0367	-3.3253	-6.4672
w_{it}^*	1.5506	-0.2651	-4.9772	-7.2921

Note: For the IPS tests, the critical value at the 5% level is -1.81 for model with an intercept and -2.44 for model with intercept and linear time trend. Individual lag lengths are based on Akaike Information Criteria (AIC).

B.2.2. Panel cointegration tests

Table B.2: Pedroni tests for different countries regimes

	Inflation		Openness		Exchange rate volatility	
	Low	High	Low	High	Low	High
<i>panel v-stat</i>	5.847	3.812	6.145	3.406	5.160	4.715
<i>panel rho-stat</i>	-5.339	-3.273	-5.297	-3.289	-3.669	-5.564
<i>panel pp-stat</i>	-5.746	-3.400	-5.527	-3.640	-4.186	-5.402
<i>panel adf-stat</i>	-5.114	-1.770	-5.608	-0.721	-2.692	-4.489
<i>group rho-stat</i>	-4.060	-3.229	-4.578	-2.509	-2.770	-4.911
<i>group pp-stat</i>	-5.520	-3.704	-5.641	-3.513	-4.080	-5.582
<i>group adf-stat</i>	-4.554	-1.832	-5.803	0.015	-2.702	-4.247

Note: Except the *v-stat*, all test statistics have a critical value of 1.64 (if the test statistic is less than 1.64, we reject the null of no cointegration). The *v-stat* has a critical value of 1.64 (if the test statistic is greater than 1.64, we reject the null of no cointegration).

PART II

PASS-THROUGH TO CONSUMER PRICES

Chapter 3

Pass-Through of Exchange Rate Shocks to Consumer Prices: Evidence from Cointegration Analysis and Pricing Chain Model

1. Introduction

After focusing on the “first-stage pass-through”, i.e. the sensitivity of import prices to changes in exchange rate movements, in the first two chapters, it is important to examine the overall effect of exchange rate changes on consumer prices, an issue which is most relevant for monetary policy in the euro area. Movements in the exchange rate can have a significant influence on inflation dynamics, both in terms of their direct effect on prices and their indirect effect through changes in the aggregate demand and wages. Thus, thorough knowledge of the underlying behavior behind pass-through is a key input to determine the proper monetary policy responses. Policymakers must be able to prevent the changes in relative prices (such as those stemming from exchange rate movements) to fuel a continuous inflationary process.

As is well-known, the Exchange Rate Pass-Through (ERPT) to consumer prices involves both first and second-stage pass-through at once, i.e. the transmission of exchange rate changes to import prices, and in turn, the transmission of import prices changes to consumer prices. Thereby, estimating the ERPT to consumer prices would include the effect of exchange-rate movements on both import prices and on other prices in the consumer basket, such as those of domestically-produced goods, services and other non-tradable prices. In order to provide reliable estimates, we need to build a framework that includes different kinds of price indices as well as the nominal exchange rate, allowing us to measure the extent of pass-through at different levels. To achieve this, MCCARTHY (2007) propose a VAR analysis that include all stages of the distribution chain (import, producer and consumer prices) to analyze how exchange rate fluctuations “pass-through” the production process from the import of products to the consumer level. Contrary to the single-equation method, this framework allows for underlying dynamic interrelations among prices at different stages of distribution and other variables of interest. The advantage of simultaneous equation approach allows for potential and highly likely endogeneity between the variables of interest, ignoring such simultaneity would result in simultaneous equation bias. In a single-equation pass-through regression, for example, the fact that domestic inflation may affect the exchange rate is ignored.

Recently, many empirical studies has adopted the modelling strategy of MCCARTHY (2007) to estimate the ERPT along the distribution chain (see e.g. CA’ZORZI, KAHN, and SÁNCHEZ, 2007; CHOUDHRI, FARUQUEE, and HAKURA, 2005; FARUQUEE, 2006; HAHN, 2003, to name but a few). However, an important drawback regarding this literature, including MCCARTHY (2007), is that the time-series properties of the data - particularly non-stationarity and cointegration issues - was neglected. To our knowledge, the only exception was HÜFNER and SCHRÖDER (2002) who estimate a Vector Error Correction Models (VECM) incorporating the long-run relationships among the variables. Deriving impulse responses functions from the VECM, the authors examine how external shocks are propagated from one price stage to the next. By contrast, HÜFNER and SCHRÖDER (2002) have not analyzed the information contained in “levels” variables, i.e. the long-run equilibrium relationship in the cointegrating vectors. In other words, they did not measure the long-run ERPT in the cointegrating relationship in addition to the impulse response analysis. On the other hand, for the

US economy, KIM (1998) has focused on the long-run relationship contained in the cointegrating space to estimate the degree of pass-through. Nevertheless, unlike the study of HÜFNER and SCHRÖDER (2002), KIM (1998) did not carry out impulse response functions analysis which could be a natural progression from the cointegration analysis. Consequently, in our empirical work, we propose to measure the long-run responsiveness of consumer prices to exchange rate depreciation, as in KIM (1998), and to derive impulse responses functions from VECM system using a pricing chain framework as in HÜFNER and SCHRÖDER (2002). To achieve this, we propose a Cointegrated VAR (CVAR) model as it allows us to take proper account of the non-stationarity of the data, i.e. look for cointegration properties in the data, and at the same time disentangle short- and long-run dynamics. This exercise is conducted for 12 euro area (EA) countries. As major problem for an analyzing pass-through in the EA is the lack of sufficiently long time series (see HÜFNER and SCHRÖDER (2002) and HAHN (2003)), our study propose a larger sample period covering the pre- and post-euro episodes.

THE OBJECTIVE OF this chapter is twofold: On one hand, we seek to remedy some of the shortcoming of the previous studies, by taking into account the non-stationarity and the endogeneity of the variables within a CVAR framework. That way, we can analyze the long-run ERPT relationship contained in the cointegrating space. In this exercise, we use a basic CVAR model to focus solely on the ERPT to consumer prices. This provides new up-to-date estimates of pass-through for the economies of the euro zone. On the other hand, in the spirit of MCCARTHY (2007), we propose an extended CVAR model that permits to track pass-through from exchange rate fluctuations to each stage of the distribution chain. The methodology of MCCARTHY (2007) is applied here with some modifications. First, the long-run proprieties of the data are considered through a Vector Error Correction Models. Second, a measure of foreign costs is included in the system as an exogenous variable which is considered as a primary variables throughout ERPT literature. We pretend that this give will us a more reliable estimates of pass-through. After estimating our *CVAR pricing model*, several analytical tools are used to explore the impact of exchange rate shocks: First, impulse responses are computed to analyze the size and speed of the pass-through of external shocks along the distribution chain; second, variance decompositions are provided to capture the relative

importance of external shocks in explaining fluctuations in the different price indices. Next, the existence of a decline in the response of consumer prices is checked; and finally, historical decompositions are used to assess how the external factors - exchange rate and import prices shocks - has contributed to the consumer inflation in the pre- and post-euro episodes.

The rest of the **Chapter 3** is organized as follows: the next section provides an overview of some VAR studies on ERPT focusing on EA countries. Section 3 outlines the baseline model used for the empirical analysis. In section 4, the data set and their properties are discussed. Section 5 contains the main results from the cointegration analysis. In section 6, we consider provide the results from the CVAR pricing chain model. Section 7 concludes the chapter 3.

2. ERPT in EA countries: Overview of VAR studies

There has been a growing interest in examining the extent of pass-through in EA countries during the last decade, although the number of studies is still relatively limited compared to empirical literature on US economy.¹ In this section we intend to give some insight on the empirical literature that used VAR models to measure the degree of ERPT in EA countries. In Table 3.1, we provide an overview of VAR studies that cover EA countries. Mainly, we emphasize on three points regarding this literature: First, the data frequency and variables employed in the VAR system. Second, type of VAR model (level, first-difference, cointegrated), techniques and tools of VAR models (impulse response functions, variance decompositions, historical decompositions) and identification schemes of the structural shocks (short-run Choleski decompositions, long-run Blanchard-Quah restrictions, both short- and long-run identifying restrictions as in GALI (1992)). Finally, the size as well as the speed of the response of consumer prices to exchange rate shock - that is, 1% currency depreciation shocks).

Among the most cited VAR study is the influential paper of MCCARTHY (2007) who investigates the pass-through on the aggregate level for selected industrialized

¹European ERPT studies have been scarce given that the time horizon since the introduction of the euro is rather short.

countries, including four EA countries, namely Belgium, Germany, France and Netherlands. The author estimates a first difference VAR model at different stages along the distribution chain, i.e. import prices, producer prices and consumer prices. In this study ERPT to consumer prices is found to be modest in most of the analyzed countries, with the exception of Belgium and Netherlands. Also, the results show that import share of a country and the persistence of exchange rate changes are found to be positively correlated with the extent of pass-through to consumer prices, while exchange rate volatility is found to be negatively correlated. A similar pricing chain model was estimated for the EA by HAHN (2003). In spite of the weakness of the ERPT, the author argued that external shocks - oil prices and exchange rate shocks together - seem to have contributed largely to inflation in the euro area since the start of the monetary union.²

Main criticism addressed to these studies is that they neglect of the time-series properties of the data, particularly non-stationarity and cointegration issues. HÜFNER and SCHRÖDER (2002) found that the endogenous variables in their VAR system are cointegrated using the Johansen procedure. Thereby, they propose to analyze the ERPT to consumer prices in the five largest countries of the EA by applying a Vector Error Correction Model (VECM) that retains information attained from any cointegrating relationships found. After aggregating the national results, the authors found a rather modest pass-through for the whole EA: four percent after one year, which rises to its long-run level of eight percent after about three years.³ Those results are obtained from the impulse responses functions which are derived from the VECM. However, HÜFNER and SCHRÖDER (2002) have not analyzed the information contained in “levels” variables, i.e. the long-run equilibrium relationship in the cointegrating vectors. In other words, they did not measure the long-run ERPT in the cointegrating relationship in addition to the impulse response analysis.

Besides, KIM (1998) estimates exchange rate pass-through for the US economy using cointegration analysis. He found that his five macroeconomic variables - producer price index, the trade weighted effective exchange rate, money supply, aggregate income

²The euro had depreciated by roughly 25 percent against the U.S. dollar in the first two years of his existence.

³Approximations for the euro area data are derived using the relative weights of each country’s inflation rate in the Harmonized Index of Consumer Prices (HICP). In table 3.1, we provide the individual pass-through estimates for the five EA countries.

and interest rates - are cointegrated. The authors estimated the long-run ERPT contained in the cointegrating space and found a significant elasticity equal to -0.24 percent following 1 percent appreciation of US dollar. Nevertheless, unlike the study of HÜFNER and SCHRÖDER (2002), KIM (1998) did not carry out impulse response functions analysis which could be a natural progression from the cointegration analysis. In our empirical work, we follow the cointegration analysis approach, as in KIM (1998), to measure the long-run responsiveness of consumer prices to exchange rate depreciation. This will be completed by impulse responses functions analysis derived from the VECM using the pricing chain framework as in MCCARTHY (2007).

It should be noted that most VAR studies on ERPT has adopted standard recursive identifying restrictions. This implies that the identified shocks contemporaneously affect their corresponding variables and those variables that are ordered at a later stage, but have no impact on those that are ordered before. It is well-known that the results derived from VAR models may strongly depend on the ordering of the variables.⁴ HAHN (2003) has carried out different identification schemes to check the robustness of the pass-through estimates. Different plausible orderings of the variables in the Choleski decomposition as well as an identification scheme that includes both short and long run restrictions was used. The author argued that these had minimal effects on the results were. Alternatively, SHAMBAUGH (2008) propose to use the BLANCHARD and QUAH (1989) methodology imposing the restriction that certain shocks cannot affect the level of certain variables in the long run. This leaves the short-run reactions free and enforces the long run assumptions to identify the shocks. Finally, MIHAILOV (2008) proposes *generalized* impulse response analysis, in the spirit of PESARAN and SHIN (1998), as an alternative to the traditional *orthogonalized* recursive one. Its main advantage that it does not require orthogonalization of shocks and, thus, it is invariant to the ordering of variables. We use this approach as a complementary check of robustness in our empirical work.

⁴FAUST and LEEPER (2003) provide a strong rejection of recursive ordering procedures that assume some variables can or cannot respond to other variables in the first period of a shock. They show that if one tests a wide variety of reasonable restrictions on the relationships between the variables, the responses to shocks can vary a great deal.

Table 3.1: Main VAR Studies on EA countries

Study	Data & Variables	Methodology	Response of consumer prices to 1% currency depreciation
Hüfner and Schröder (2002)	<p>Monthly data from 1982:1 to 2001:1 for five large EA countries (France, Germany, Italy, Netherlands and Spain)</p> <p><i>Endogenous variables:</i> Oil price, NEER, output gap, interest rate and 3 price indices (import prices, producer prices and consumer prices)</p>	<p>Cointegration Analysis using Johansen procedure</p> <p>Impulse responses and variance decompositions derived from the VECM</p> <p>Identification of shocks by Cholesky decomposition</p>	<p>France: 0.01 (6 months), 0.07 (12 months), 0.12 (18 months), 0.16 (24 months)</p> <p>Germany: 0.07 (6 months), 0.08 (12 months), 0.09 (18 months), 0.10 (24 months)</p> <p>Italy: 0.06 (6 months), 0.12 (12 months), 0.16 (18 months), 0.18 (24 months)</p> <p>Netherlands: 0.12 (6 months), 0.11 (12 months), 0.11 (18 months), 0.11 (24 months)</p> <p>Spain: 0.09 (6 months), 0.08 (12 months), 0.08 (18 months), 0.08 (24 months)</p>
Hahn (2003)	<p>Quarterly data from 1970:2 to 2002:2 for the euro area</p> <p><i>Endogenous variables:</i> Oil prices, interest rate, output gap, exchange rate, non-oil import prices, producer prices and HICP</p>	<p>Impulse responses, variance and historical decompositions derived from a first difference VAR model</p> <p>Identification of shocks by Cholesky decomposition</p>	<p>1st quarter: 0.025</p> <p>1st year: 0.08</p> <p>3 years: 0.16</p>
Choudhri <i>et al.</i> (2005)	<p>Quarterly series at annual rates 1979:1 to 2001:3 for non-US G-7 countries</p> <p><i>7 endogenous variables:</i> Interest rate, exchange rate, import price, export price, producer price, consumer price and wage rate</p> <p><i>2 exogenous variables:</i> Foreign interest and foreign consumer price</p>	<p>Impulse responses derived from restricted VAR model</p> <p>Identification of shocks using structural short-run restrictions</p>	<p>Germany: 0.15 (1 quarter), 0.20 (4 quarters), 0.36 (10 quarters)</p> <p>France: 0.00 (1 quarters), 0.10 (4 quarters), 0.09 (10 quarters)</p> <p>Italy: 0.02 (1 quarter), 0.14 (4 quarters), 0.26 (10 quarters)</p>

Table 3.1: *Continued*

Study	Data & Variables	Methodology	Response of consumer prices to 1% currency depreciation
Faruqee (2006)	Monthly data from 1990 to 2002 for the euro area <i>Endogenous variables:</i> Nominal effective exchange rate, wages, import prices, export prices, producer prices, consumer prices	Impulse responses derived from VAR in first differences Identification of shocks by Cholesky decomposition	0.00 after 1 month 0.01 after 6 months 0.02 after 12 months 0.02 after 18 months
McCarthy (2007)	Quarterly data from 1976:1 to 1998:4 for 9 developed countries among them 4 EA members (Germany, France, Belgium and Netherlands) <i>Endogenous variables:</i> Oil price, NEER, output gap, import prices, producer prices, consumer prices, interest rate and monetary aggregate	Impulse responses, variance and historical decompositions derived from a first difference VAR model Identification of shocks by Cholesky decomposition	ERPT is particularly large in Belgium and Netherlands. Wrong (negative) sign for France. By the end of two years the response is imprecisely estimated.
Shambaugh (2008)	Quarterly data from 1973:1 to 1999:4 for 16 countries among them 4 EA members (Austria, Finland, Germany and Greece) <i>Endogenous variables:</i> Industrial production, real exchange rate, CPI, nominal exchange rate, import price	Impulse responses and variance decompositions derived from a first difference VAR model Blanchard-Quah long run restrictions methodology	<i>ERPT ratio following external shock:</i> Austria: 0.83 (1 quarter), 0.55 (4 quarters) Finland: 0.71 (1 quarter), 0.79 (4 quarters) Germany: 0.32 (1 quarter), 0.37 (4 quarters) Greece: 0.25 (1 quarter), 0.70 (4 quarters)

3. Empirical Methodology

Initially, our analysis aims at capturing the effects of changes in exchange rates on consumer prices which is the key variable for the policy issues. Thus, we start with a baseline model, in the spirit of KIM (1998), that relates consumer prices (cpi_t) to the trade weighted effective exchange rate (e_t), oil price (oil_t), aggregate income (y_t) and interest rates (r_t) in cointegrated VAR (CVAR) framework. Using Johansen procedure, CVAR analysis could be useful in this context as it allows us to take proper account of the non-stationarity of the data, looking for cointegration properties in the data, and at the same time disentangle short- and long-run dynamics. Thus, it enables retention of the important information contained in levels variables. This levels information is lost in more traditional first-difference VAR models.

As a starting point of the analysis, we consider the following vector of endogenous variables:

$$x_t = (cpi_t, e_t, oil_t, y_t, r_t)' \quad (3.1)$$

Having firstly tested the stationarity of the variables, we apply cointegration tests for each country to check whether long-term relationships exist between the variables. The Johansen test is used to assess whether or not cointegration exists in the system of variables. In order to describe this, we begin firstly by considering the following system of five-equation VAR(k) model:

$$x_t = A_1 x_{t-1} + \dots + A_k x_{t-k} + \mu + \psi D_t + \varepsilon_t, \quad t = 1, 2, \dots, T, \quad (3.2)$$

Equation (3.2) can be expressed as an error or vector equilibrium correction model (VECM), i.e. a CVAR, which is formulated in terms of differences as follows:

$$\Delta x_t = \Gamma_1 \Delta x_{t-1} + \dots + \Gamma_{k-1} \Delta x_{t-k+1} + \Pi x_{t-1} + \mu + \psi D_t + \varepsilon_t \quad (3.3)$$

Where x_t is a (5×1) vector of $I(1)$ endogenous variables as given in Equation (3.1); k is lag length; μ is a constant term; D_t is a vector including deterministic variables (centered seasonal dummies and intervention dummies) and weakly exogenous variables; and ε_t is a $(k \times 1)$ vector of errors which are assumed identically and independently distributed and follow a Gaussian distribution $\varepsilon_t \sim \text{iid } N_p(0, \Omega)$, with Ω denotes the variance-covariance matrix of the disturbances.

The VECM representation, i.e. the CVAR model, encompasses both short- and long-run information of the data. The matrix Π assembles the long-run information and the Γ_i s contain the short-run properties. $\Pi = \alpha\beta'$ has reduced rank r . The matrices α and β are of dimension $(5 \times r)$, α depicts the speed of adjustment, and β represents the cointegrating vectors. The Johansen procedure estimates equation (3.3) subject to the hypothesis that Π has a reduced rank $r < 5$. This hypothesis can be written as:

$$H(r) = \alpha\beta' \quad (3.4)$$

JOHANSEN and JUSELIUS (1990) show that, under certain circumstances, the reduced rank condition implies that the processes Δx_t , and $\beta' x_t$, are stationary even though x_t , itself is non-stationary. The stationary relations $\beta' x_t$, are referred to as cointegrating relations. To determine the number of cointegrating vectors (r) in the system, i.e. the cointegration rank, we employ the widely used trace test statistics (the results of which are reported in Table C.4 in Appendix C.3):

$$\text{Trace} = -N \sum_{i=r+1}^5 \ln(1 - \hat{\lambda}_i) \quad (3.5)$$

Where N is the number of observations and $\hat{\lambda}_i$ is the estimated eigenvalue. When the appropriate model has been identified for the system in terms of lag length and cointegration rank, the coefficients on the α matrix reveal the long-run dynamic while the coefficients on the β matrix reveal the drivers towards the long-run equilibrium.

In order to determine the responsiveness of consumer prices to exchange rate changes, the coefficient estimates of the cointegrating vectors are normalized on consumer prices. Thus, the coefficients on exchange rate indicate the degree of ERPT. Also, in the cointegration analysis, we focus on the first cointegrating vector. As discussed in JOHANSEN and JUSELIUS (1992), the first cointegrating vector is the most associated with the stationary part of the model since it has the highest eigenvalue. After estimating the ERPT coefficient in the long-run, we follow BEIRNE and BIJSTERBOSCH (2011) by testing a number of restrictions on the long-run parameters in order to examine specific hypotheses on pass-through:

H₁: Full ERPT to consumer prices with other long-run parameters unrestricted, i.e. test of whether the first cointegrating is as follows

$$\beta_1' = \begin{matrix} & cpi & e & oil & y & r \\ (1 & 1 & \gamma & \eta & \varphi) \sim I(0) \end{matrix}$$

H₂: Full ERPT to consumer prices with zero constraints on other long-run parameters, i.e. test of whether the first cointegrating is as follows

$$\beta_1' = \begin{matrix} & cpi & e & oil & y & r \\ (1 & 1 & 0 & 0 & 0) \sim I(0) \end{matrix}$$

H₃: Zero ERPT to consumer prices with other long-run parameters unrestricted, i.e. test of whether the first cointegrating is as follows

$$\beta_1' = \begin{matrix} & cpi & e & oil & y & r \\ (1 & 0 & \gamma & \eta & \varphi) \sim I(0) \end{matrix}$$

H₄: Zero ERPT to consumer prices with zero constraints on other long-run parameters, i.e. test of whether the first cointegrating is as follows

$$\beta_1' = \begin{matrix} & cpi & e & oil & y & r \\ (1 & 0 & 0 & 0 & 0) \sim I(0) \end{matrix}$$

If \mathbf{H}_1 or \mathbf{H}_2 holds, this would imply that exchange rate changes are fully transmitted to consumer prices, while if \mathbf{H}_3 or \mathbf{H}_4 holds, there is a null pass-through, i.e. consumer prices do not respond to currency movements.

After achieving cointegration analysis, we carry out an impulse response function analysis on the VECM that includes the distribution chain of pricing (import prices, producer prices and consumer prices) as in MCCARTHY (2007) and HÜFNER and SCHRÖDER (2002). This enables us to assess both the size as well as the speed of the pass-through in the EA countries. This is performed using the traditional orthogonalized Cholesky decomposition to identify the structural shocks. In fact, the recursive structure embodied in the approach, implying that the variables in the system do not react contemporaneously to shocks imposed on variables ordered later, means that it is important to ensure a correct ordering scheme. To ensure the robustness of our results, we also compute the generalized impulse response functions as proposed by PESARAN and SHIN (1998) where ordering of the variables does not matter. Besides, variance decompositions are computed to capture the relative importance of external shocks (exchange rate and import prices shocks) in explaining fluctuations in consumer prices. Furthermore historical decompositions are will be employed to assess the role and the importance of external shocks on inflation in the euro area during different episodes.

4. Data selection and their properties

In order to measure the effects of exchange rate changes on consumer prices, we start with a baseline VECM that contains five endogenous variables. In addition to our two key variables - exchange rate and consumer prices - we have included three macroeconomic variables affect the inflation of consumer prices directly. The choice of the variables is based on the following considerations: first, oil prices enter the VECM to controls for the impact of supply shocks; second, to balance the model with respect to the demand side, a measure of national income is added in the baseline model; and finally, a short-run interest rate is included to allow for the effects of monetary policy.⁵ As discussed by PARSLEY and POPPER (1998), taking into account monetary

⁵With the exception of interest rates, all variables are in logs.

policy significantly improves the estimation results of ERPT. In fact, central banks are concerned with keeping domestic inflation within its target range which may insulate prices from exchange rate movements. Thus, neglecting the effects of monetary policy results in the common omitted variables problem. In a subsequent step, our basic model will be augmented to include the whole pricing, i.e. import prices, producer prices and consumer prices.

In this study, we focus our analysis on 12 EA countries ((Austria, Belgium, Germany, Spain, Finland, France, Greece, Ireland, Italy, Luxembourg, Netherlands and Portugal). For each country a set of quarterly data was collected covering the time period 1980:1 to 2010:4. The consumer price (cpi_t) is the overall consumer price index to provide the broadest measure of inflation at the consumer level. We did not use the Harmonized Index of Consumer Prices (HICP) due the short data availability of this variable. Exchange rate data are effective nominal exchange rates of the national currencies which use the trade weights of each country.⁶ The oil price (oil_t) is represented by a crude oil price index denominated in US dollar in order to avoid multicollinearity issues with the exchange rate.⁷ The national income (y) is proxied by the real GDP. The 3-month interest rate is used to model monetary policy. When considering the pricing chain model (in section 6), we include the non-oil import prices as a measure of import prices (mpi_t) to avoid double-counting with oil prices index and the producer prices index (ppi_t) in manufacturing. To collect data, we have followed a cascade order, choosing when possible only one institutional source, i.e. IMF's *International Financial Statistics* and OECD's *Main Economic Indicators* and *Economic Outlook*, in that order.

Next, we check the non-stationarity of the data. In order to test this, each of the variables are tested for unit roots using the traditional ADF-test which tests the null hypothesis of non-stationarity. To ensure robustness the order of integration of the variables, ADF test is supplemented by two stationarity test. First, the Kwiatkowski-Phillips-Schmidt-Shin (KPSS) test which is structured under the opposite null hypothesis that of stationarity against a unit root alternative. Second, the DF-GLS test, proposed by

⁶The nominal effective exchange rate is defined as domestic currency units per unit of foreign currencies, which implies that an increase represents a depreciation for domestic country.

⁷McCarthy uses local price of oil to identify supply shocks, but this will include the exchange rate effect. Thus, much of the exchange rate effect may be mixed into the supply shock.

ELLIOTT, ROTHENBERG, and STOCK (1996), which is an augmented Dickey-Fuller test, similar to the test performed Dickey-Fuller tests, except that the time series is transformed via a generalized least squares (GLS) regression before performing the test. ELLIOTT, ROTHENBERG, and STOCK (1996) have shown that this test has significantly greater power than the previous versions of the augmented Dickey-Fuller test. In constructing the unit root tests, the variables in levels were tested in the presence of both an intercept and trend. The subsequent tests of first differences included only an intercept given the lack of trending behaviour in the first-differences series.

Results of the unit root tests of the variables reveal that the majority of the variables to have been generated via an integrated of order one $I(1)$ process (see Table C.1 in the Appendix C.1). First-differences variables are found to be stationary in at least two of the three tests undertaken for most cases. We can summarize the results of the three unit root tests as follows: According to ADF tests all variables are stationary in first differences with exception of consumer prices in Ireland; for the KPSS test, the null hypothesis of stationarity is accepted for most of the variables in first differences except for consumer prices in Netherlands and Portugal, while import prices are stationary in level for Luxembourg. Finally, we find that all variables $I(1)$ within DF-GLS test, with the exception of consumer prices for Luxembourg and Portugal, nominal effective exchange rate for Portugal and producer prices in Ireland.

Building on these results, the Johansen cointegration tests were undertaken to assess the existence of long-run equilibrium relationships among the variables. Given that the choice of the rank of Π should be made on the basis of a well-specified model, it is important to include the appropriate number of lags before rank tests are undertaken. Thus, the lag structure for each VECM was based on assessment of the AIC compatible with well-behaved residuals. Results from trace test, reported in Table C.4 in Appendix C.3, indicates the presence of one cointegrating vectors at least for each EA country (as in Austria and Netherlands). The null hypothesis of no cointegration was rejected for all our EA countries, with a cointegration rank identified of between one and three. A summary of the number of cointegrating equations (CE) identified across each country as well as the optimal lag length is reported in Table 3.2.

5. Cointegration Analysis

In the section, we focus on the long-term part of our baseline VECM, i.e. the long-run relationships present in the cointegrating space. Our primary concern is to assess the relative signs and magnitudes of the long-run ERPT coefficients across EA countries. To this end, there are some issues that must be considered here. First, ERPT equation must contain a proxy for foreign costs as recommended by the bulk of empirical literature (see GOLDBERG and KNETTER (1997)). Given that foreign costs are an exogenously determined variables regarding our EA countries, we propose to include a proxy for costs of a country's trading partners as an exogenous variables in our basic VECM. Therefore, to capture changes in foreign costs, we construct a typical export partners cost proxy (W_t^*) that used throughout the ERPT literature (see *inter alia* BAILLIU and FUJII (2004) and CAMPA and GOLDBERG (2005)): $W_t^* = Q_t \times W_t/E_t$, where Q_t is the unit labor cost based real effective exchange rate, W_{it} is the domestic unit labor cost and E_t is the nominal effective exchange rate. Taking the logarithm we obtain the following expression: $w_{it}^* = q_t + w_t - e_t$. Since the nominal and real effective exchange rate series are trade weighted, we obtain a measure of foreign firms' costs with each partner weighted by its importance in the domestic country's trade.⁸ Second, besides the seasonal dummy variables a shift, we introduce dummy in 1990:07 (D_{90}) and kicks in until the end of the sample. Chow tests for multivariate models, as introduced by CANDELON and LUTKEPOHL (2001), denote the presence of structural break in vicinity of 1990 (see Table C.5 in Appendix C.5).⁹ Including D_{90} helps to restore the stability of the cointegrating vectors. In Table C.1 in Appendix C.4, we investigate parameter constancy by means of recursive estimates of the eigenvalues. Plots reveals that recursive estimates of the eigenvalues, over a 40-month window, are broadly constant for most of EA countries which is an indication of the stability of the cointegrating vectors identified. It is worth noting that centered seasonal dummies, shift dummy and exogenous foreign

⁸To measure the extent of pass-through in the non-US G-7 countries, CHOUDHRI, FARUQUEE, and HAKURA (2005) enter two foreign exogenous variables - foreign interest rate and the foreign consumer price index - in their first-difference VAR model.

⁹We can select May 1998, the month on which the parities among European currencies replaced by the euro were announced, as the date for the break. However, as showed in Chapter 1 and in most of empirical literature (see CAMPA and GOLDBERG (2002, 2005) among others), the date of creation of the euro has not been found as a regime shifts in the monetary union countries.

costs enter the vector D_t in equation (3.3). Final issue concerns the specification of VECM of each of our 12 EA countries. In most of the cases the most appropriate model appears to be that which includes a trend in the cointegrating equation and permits the intercept to enter both the cointegration space and the VAR, i.e. unrestricted intercept and restricted trend. The only exceptions are Spain, Ireland and Luxembourg where we include only a constant in the cointegrating equations and in the short-term part of the VECM, i.e. unrestricted intercept.¹⁰ Summary of our 12 CVAR models are reported in Table 3.2.

Table 3.2: Summary of basic CVAR Models

Country	VAR lags	Number of CE	Model specification
Austria	2	1	Restricted trend
Belgium	2	2	Restricted trend
Germany	2	3	Restricted trend
Spain	3	2	Unrestricted intercept
Finland	1	2	Restricted trend
France	2	3	Restricted trend
Greece	5	2	Restricted trend
Ireland	3	3	Unrestricted intercept
Italy	2	3	Restricted trend
Luxembourg	3	2	Unrestricted intercept
Netherlands	2	1	Restricted trend
Portugal	2	3	Restricted trend

Note: The optimal number of lags in the VECM was determined using the AIC criterion. The number of cointegrating equations is equal to the number of cointegration equations found by the Johansen trace test.

5.1. Long-run ERPT to consumer prices

As we mentioned above, we focus on the first (most statistically significant) cointegrating equation to measure the extent of pass-through in the long-run. The long-run parameters for each unrestricted CVAR model are reported in Table 3.3. In unrestricted form, it is clear that the signs of the parameters appear in most cases to accord with priors. In most of case, positive coefficients are observed on exchange rate, oil

¹⁰The use of unrestricted intercepts and restricted trends is consistent with data that exhibit some form of trending behaviour. When we expect some of the data to be trend stationary, a good idea is to start with a restricted linear trend and then test the significance of the trends.

prices, real GDP and the interest rate series.¹¹ Thereby, a rise in exchange rate (i.e. depreciation), in oil prices, in real GDP or in interest rate is associated with a higher domestic consumer prices. In some cases, there appears to be some inconsistency regarding the sign on GDP or interest rate, but roughly speaking, our results tend to agree with the expected signs.

Table 3.3: Coefficients of first cointegrating vector

Country	cpi_t	e_t	oil_t	gdp_t	r_t	$trend$
Austria	1,000	0,248* (1,799)	0,124** (2,494)	0,712*** (10,542)	-0,038*** (-10,594)	0,026* (1,819)
Belgium	1,000	0,282*** (3,800)	0,468*** (3,373)	-0,213 (-0,506)	0,019*** (4,250)	0,007*** (2,835)
Germany	1,000	0,169** (2,305)	0,464 (1,523)	0,968*** (2,606)	0,073*** (8,608)	0,011*** (6,302)
Spain	1,000	0,337** (2,254)	0,535* (1,910)	0,880 (1,538)	0,002 (0,310)	-
Finland	1,000	0,117* (1,897)	0,413*** (2,915)	0,578*** (7,345)	-0,009*** (-2,946)	-0,004* (-1,753)
France	1,000	0,166*** (2,693)	0,279** (2,260)	-0,290 (-0,752)	0,013*** (2,781)	0,006*** (2,747)
Greece	1,000	0,576*** (4,494)	1,027*** (5,416)	0,371 (1,002)	-0,036*** (-3,957)	0,031*** (6,192)
Ireland	1,000	0,397*** (4,009)	0,208** (2,495)	0,485*** (5,126)	0,003 (1,571)	-
Italy	1,000	0,352*** (5,231)	0,486*** (3,394)	1,098*** (3,129)	0,012*** (2,720)	0,003* (1,813)
Luxembourg	1,000	0,339*** (5,472)	0,667*** (4,801)	0,468*** (13,338)	0,008** (2,089)	-
Netherlands	1,000	0,298*** (4,400)	0,683*** (5,327)	-0,637*** (-13,048)	-0,039*** (-10,820)	0,044*** (8,236)
Portugal	1,000	0,833*** (8,553)	0,056 (1,244)	-0,084 (-0,206)	0,013* (1,739)	0,018*** (2,732)

Note: *, ** and *** denote significance level at 10%, 5% and 1% respectively. *t*-stat are in parentheses.

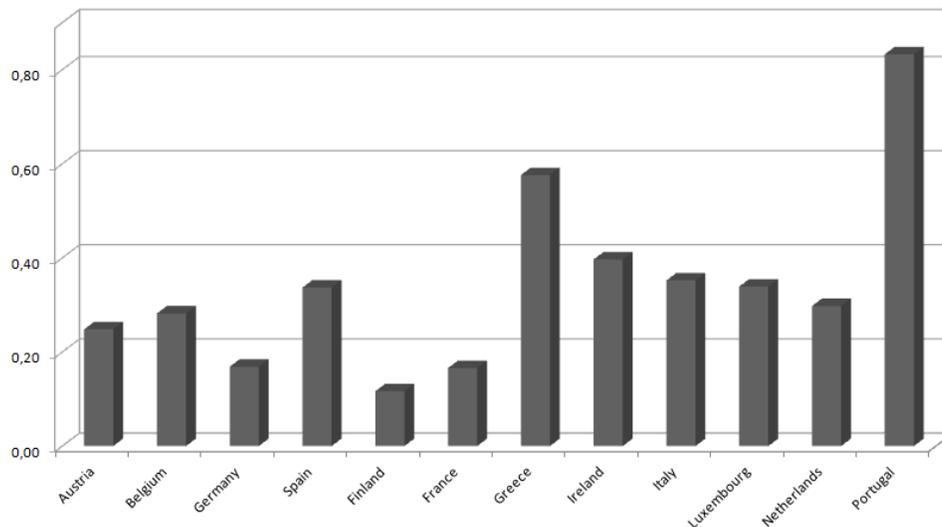
Concerning the degree of ERPT, our results point out cross-country differences in the responsiveness of consumer prices in the long-run (see Figure 3.1).¹² Germany, Finland and France have the lowest coefficients in our sample of EA, with long-run ERPT not exceeding 0.20%. The degree of ERPT appears to be most prevalent in

¹¹The positive relationship between consumer prices and interest rate is consistent with the long-run Fisher effect.

¹²All long-run rates of pass-through are significantly different from zero in our sample of EA countries.

Portugal and Greece. For Portugal, a 1% depreciation of exchange rate increases domestic consumer prices by roughly 0.84%, while for Greece, consumer prices rise by 58% following one percent depreciation of exchange rate. In their study on 20 industrialized countries, GAGNON and IHRIG (2004) found that Portugal and Greece have the highest long-run response of consumer prices over the period 1972 to 2000. The pass-through elasticities are: 0.43 percent for Portugal and 0.52 percent for Greece. Nevertheless, these pass-through coefficients are still lower compared to our results. As a matter of fact, GAGNON and IHRIG (2004) did not find any evidence of cointegration between variables in levels, that's why they estimate their pass-through *single-equation* in first-differences. Thus, their definition of "long-run" effect stems from the feedback effects resulting from the inclusion of the lagged dependent inflation terms (dynamic equation).¹³ We see that taking into account the times series proprieties of the date (non-stationarity and cointegration relationship) may give a more reliable long-run ERPT estimates.

Figure 3.1: Long-run ERPT in EA countries



Source: Personal calculation.

¹³The effects of an exchange rate change in period t will influence inflation over several periods subsequent to this as a result of these feedback effects.

Moreover, it should be noted that the response of consumer prices is still weak in comparison to import prices (see our results in Chapter 1 and Chapter 2). Several explanations have been put forward by ERPT literature. In fact, imported goods have to go through distribution sector before they reach consumers in domestic country. Thus, local distribution costs (such as transportation costs, marketing, and services), may cause a wedge between import and consumer prices. Also, competitive pressure in distribution sectors may explain why consumer prices do not respond dramatically to exchange rate changes. As discussed in BACCHETTA and VAN WINCOOP (2002), the weakness of CPI inflation reaction to exchange rate changes is due, in part, to differences in the optimal pricing strategies of foreign producers and domestic wholesalers/retailers. Due to competitive pressure in the domestic market, domestic wholesalers import goods priced in foreign currency (PCP) and resell them in domestic currency (LCP). This would entail much lower ERPT to CPI inflation than expected. Finally, we can add that substitution effect can occur. If home currency is depreciating, domestic firms or wholesalers may reduce sourcing foreign products (since their price becomes higher), shifting towards substitute domestically produced goods. That way, consumer prices would be more insulated from exchange rate movements. In the next section 6, we give more insights on how exchange rate effect declines along the distribution chain. Referring to the adjustment speed, in the presence of price stickiness, adjustment lags at different stages of distribution might accumulate. This would explain the lack of response of consumer prices compared to imports prices.

5.2. Speed of adjustment to long-run equilibrium

In Table 3.4, we set out the adjustment coefficients (or loading factors) revealing the speed with which the long-run equilibrium is achieved. It is known that lack of significance on these parameters indicates the presence of weak exogeneity, meaning that the variable does not respond to or correct for deviations to the long-run equilibrium. For oil prices, we find non-significant adjustment coefficients in the half of EA countries, this could be a sign of weak exogeneity. We could impose weak exogeneity on oil prices but it does not alter the long-run coefficients of ERPT. We keep on only foreign as the only exogenous variables in our CVAR models.

Table 3.4: Adjustment coefficients in the basic VECM

	Austria	Belgium	Germany	Spain	Finland	France
cpi_t	0,018*** (8,071)	-0,056*** (-8,626)	-0,026*** (-7,733)	-0,053*** (-6,782)	-0,013*** (-2,514)	-0,042*** (-8,395)
e_t	0,069*** (3,050)	-0,083 (-1,022)	-0,075* (-1,689)	-0,153*** (-2,648)	0,106*** (4,162)	-0,039* (-1,621)
oil_t	-0,096 (-1,451)	0,018** (2,447)	0,041* (1,786)	0,024 (0,783)	0,006 (0,171)	0,069*** (3,473)
gdp_t	-0,003** (-1,998)	0,016*** (2,573)	0,049*** (3,386)	-0,021** (-2,291)	-0,009* (-1,641)	0,004 (0,531)
r_t	-0,453* (-1,630)	1,822* (1,727)	0,875** (2,370)	0,552 (0,382)	1,666* (1,873)	-0,047*** (-3,286)
	Greece	Ireland	Italy	Luxembourg	Netherlands	Portugal
cpi_t	0,033** (2,342)	-0,119*** (-11,256)	-0,044*** (-6,860)	-0,073*** (-6,110)	0,022*** (6,313)	-0,041*** (-7,424)
e_t	0,238*** (6,179)	-0,225*** (-2,723)	-0,027 (-0,304)	-0,070** (-2,516)	0,173*** (3,895)	-0,058** (-2,156)
oil_t	-0,027* (-1,694)	0,007** (-2,127)	0,000 (-0,010)	-0,019 (-0,289)	-0,047** (-2,047)	0,010 (0,755)
gdp_t	-0,016 (-0,498)	0,058*** (-2,709)	-0,002 (-0,178)	-0,116** (-2,295)	-0,020** (-2,550)	-0,009** (-2,448)
r_t	0,041** (2,049)	-7,239* (-1,849)	1,306 (0,857)	-10,805*** (-5,393)	-0,531 (-1,043)	1,576*** (3,303)

Note: *, ** and *** denote significance level at 10%, 5% and 1% respectively. t -stat are in parentheses.

Also, it is important to assess the dynamics of adjustment to the long-run equilibrium on consumer prices equations. For example, in France, the error correction mechanism containing cpi_t enters its own equation with a coefficient of adjustment equal to -0,042 and a highly significant t -statistic of -8,395. This means that when consumer prices exceed their long-run equilibrium level, they adjust downwards at a rate of 4.2% per quarter until equilibrium is restored. This implies a long period of half-life adjustment which is equal to four years.¹⁴ For information, the half-life measures are calculated as follows: for consumer prices equation in France, we see that adjustment coefficient is 0.042. We know that $(1 - 0.042)^n = 0.5$, where n is the number of periods in the half-life of deviations of cpi_t from equilibrium. Taking natural logs of both sides of the equation and rearranging gives $n = (\ln 0.5 / \ln 0.96) \simeq 16$ (quarters). According to Table 3.4, a similar slow adjustment of consumer prices towards equilibrium is found

¹⁴The so-called half-life is defined as the expected time to revert half of its deviation from the long-run equilibrium.

across our sample of EA countries.¹⁵ In fact, this slow adjustment would explain why ERPT coefficients are very weak in the short-run, as reported in the literature.

5.3. Testing hypothesis

Final step in our cointegration analysis, we turn to the number of restrictions on the long-run parameters postulated in section 3. Thus, we explore the hypotheses of full ERPT (\mathbf{H}_1 and \mathbf{H}_2) and Zero ERPT (\mathbf{H}_3 and \mathbf{H}_4). Regarding the tests of restrictions on the long-run parameters in Table 3.5, it is clear that \mathbf{H}_1 is rejected for 9 out of 12 EA countries, implying that ERPT is not complete for this sub-sample. However, We cannot reject the hypothesis of full pass-through for Portugal and Greece. These findings provide corroboration for our earlier empirical results that these two countries have the highest degree of ERPT in our sample of 12 EA.

\mathbf{H}_2 is also rejected but for all countries, indicating that complete ERPT is rejected when other variables in the system (oil prices, real GDP and interest rate) are constrained to have no effect on domestic consumer prices. Concerning \mathbf{H}_3 , the hypothesis of null ERPT is rejected for all EA countries except for Austria, Finland and France. For the latter countries, the weakness of degree of pass-through was confirmed throughout the empirical literature. For instance, GAGNON and IHRIG (2004) found the lowest ERPT elasticity in Finland with a coefficient equal to 0.01%. Finally, the hypothesis of zero ERPT when other variables in the system are constrained to have no effect on consumer prices, namely \mathbf{H}_4 , is rejected for the whole of our EA sample.

¹⁵The faster adjustment is found in Ireland with a half-life measure of one and a half years.

Table 3.5: Restrictions on long-run parameters

Country	Full ERPT		Zero ERPT	
	H_1	H_2	H_3	H_4
Austria	8,937 [0,030]	33,623 [0,000]	1,230 [0,267]	27,519 [0,000]
Belgium	8,396 [0,004]	8,478 [0,076]	3,290 [0,070]	11,780 [0,019]
Germany	3,684 [0,055]	42,011 [0,000]	2,284 [0,103]	44,954 [0,000]
Spain	3,793 [0,051]	18,468 [0,001]	2,568 [0,109]	19,335 [0,017]
Finland	7,708 [0,005]	51,644 [0,000]	2,284 [0,131]	15,558 [0,004]
France	13,753 [0,000]	26,356 [0,000]	1,475 [0,225]	20,917 [0,001]
Greece	1,157 [0,282]	12,883 [0,012]	4,210 [0,040]	15,960 [0,003]
Ireland	3,342 [0,068]	28,287 [0,000]	2,891 [0,089]	12,041 [0,017]
Italy	10,882 [0,001]	20,073 [0,000]	5,005 [0,025]	9,750 [0,045]
Luxembourg	2,744 [0,098]	17,919 [0,001]	3,768 [0,052]	17,696 [0,001]
Netherlands	7,422 [0,006]	39,723 [0,000]	4,918 [0,027]	36,272 [0,000]
Portugal	1,524 [0,294]	8,829 [0,066]	2,593 [0,094]	20,463 [0,001]

Note: Restrictions based on Likelihood Ratio tests with a chi-squared distribution, with the number of degrees of freedom equal to the number of restrictions imposed; p -values in square brackets.

6. Evidence from Pricing Chain model

As a natural progression from the cointegration analysis, we carry out impulse response functions analysis to estimate the pass-through effect of changes in the effective exchange rate to the domestic prices. Then, in this section we expand our baseline CVAR model to incorporate features of a distribution chain pricing framework in the spirit of MCCARTHY (2007) and HÜFNER and SCHRÖDER (2002). The *CVAR pricing chain model* enables us to examine the pass-through at different stages along the distribution chain, i.e. import prices, producer prices and consumer prices. This exercise is of great interest for EA price analysis as it reveals how exchange rate shocks are propagated from

one price stage to the next. Thus, a measure of import prices (mpi_t) and producer prices (ppi_t) will be introduced in our basic CVAR model to obtain the following system of variables:

$$x_t = (oil_t, y_t, e_t, mpi_t, ppi_t, cpi_t, r_t)' \quad (3.6)$$

The impulse responses functions are derived from a system of seven-equation VECM that incorporate the long-run relationships among the variables.¹⁶ This framework allows for underlying dynamic interrelations among prices at different stages of distribution and the rest of variables. It furthermore enables to trace the dynamic responses of prices to exchange rate shocks, i.e. it captures both the size as well as the speed of the pass-through. In addition to impulse responses, variance decompositions are computed to capture the relative importance of different shocks in explaining fluctuations in the different price indices. Furthermore, we use historical decompositions to examine the influence and the contribution of exchange rate and import prices shocks to consumer prices variation during two sub-sample periods: during the first and second stage of EMU; and since the creation of the euro till the end of our time sample. To this end, structural shocks in our *CVAR pricing chain model* must be identified. In the following subsection 6.1, we present the recursive identification scheme in the spirit of MCCARTHY (2007) that we apply in our empirical work.

6.1. Identification Scheme: MCCARTHY (2007) approach

In an influential paper MCCARTHY (2007) proposes the following identification scheme to identify the shocks in the pricing chain model. According to this scheme, inflation at each stage of distribution chain - import, producer, and consumer - in period t is assumed to comprise several components. The first component is the expected inflation at that stage based on the available information at the end of period $t - 1$. The second and third components are the effects of period t domestic supply and demand shocks on

¹⁶Johansen trace tests are applied on the seven-equation VECM which indicates the presence of one cointegrating vectors at least for each EA country. Also, the appropriate number of lags was calculated using AIC selection criterion (see Table C.3 in Appendix C.2).

inflation at that stage. The fourth component is the effect of exchange rate shocks on inflation at a particular stage. Next components are the effects of shocks at the previous stages of the chain. Finally, there is that stage's shock.

The inflation shocks at each stage are simply that portion of that stage's inflation which cannot be explained using information from period $t - 1$ plus information about domestic supply and demand variables, exchange rates, and period t inflation at previous stages of the distribution chain. These shocks can thus be thought of as changes in the pricing power and markups of firms at these stages. Two other features of the model are worthy of note. First, the model allows import inflation shocks to affect domestic consumer inflation both directly and indirectly through their effects on producer inflation. Second, there is no contemporaneous feedback in the model: for example, consumer inflation shocks affect inflation at the import and producer stages only through their effect on expected inflation in future periods.

Under these assumptions, the inflation rates of country i in period t at each of the three stages - import (mpi_t), producer (ppi_t), and consumer (cpi_t) - can be written as:

$$\Delta mpi_{it} = E_{t-1}(\Delta mpi_{it}) + \delta_{1i}\varepsilon_{it}^s + \delta_{2i}\varepsilon_{it}^d + \delta_{3i}\varepsilon_{it}^e + \varepsilon_{it}^{mpi} \quad (3.7)$$

$$\Delta ppi_{it} = E_{t-1}(\Delta ppi_{it}) + \varphi_{1i}\varepsilon_{it}^s + \varphi_{2i}\varepsilon_{it}^d + \varphi_{3i}\varepsilon_{it}^e + \varphi_{4i}\varepsilon_{it}^{mpi} + \varepsilon_{it}^{ppi} \quad (3.8)$$

$$\Delta cpi_{it} = E_{t-1}(\Delta cpi_{it}) + \eta_{1i}\varepsilon_{it}^s + \eta_{2i}\varepsilon_{it}^d + \eta_{3i}\varepsilon_{it}^e + \eta_{4i}\varepsilon_{it}^{mpi} + \eta_{5i}\varepsilon_{it}^{ppi} + \varepsilon_{it}^{cpi} \quad (3.9)$$

Where ε_{it}^s , ε_{it}^d and ε_{it}^e are the supply, demand, and exchange rate shocks respectively; ε_{it}^{mpi} , ε_{it}^{ppi} and ε_{it}^{cpi} are the import price, producer price, and consumer price inflation shocks; and $E_{t-1}(\cdot)$ is the expectation of a variable based on the information set at the end of period $t - 1$. The shocks are assumed to be serially uncorrelated as well as uncorrelated with one another within a period.

The structure of the model (3.7) to (3.9) is a part of a recursive VAR framework. Thus, to complete the empirical model, the following assumptions are added. First, supply shocks (ε_{it}^s) are identified from the dynamics of oil price inflation (Δoil_t) denominated in the local currency. Second, demand shocks (ε_{it}^d) are identified from

the dynamics of the GDP growth (Δy_t) in the country after taking into account the contemporaneous effect of the supply shock. Finally, exchange rate shocks (ε_{it}^e) are identified from the dynamics of exchange rate depreciation (Δe_t) after taking into account the contemporaneous effects of the supply and demand shocks.

$$\Delta oil_{it} = E_{t-1}(\Delta oil_{it}) + \alpha_{1i} \varepsilon_{it}^s \quad (3.10)$$

$$\Delta y_{it} = E_{t-1}(\Delta y_{it}) + \beta_{1i} \varepsilon_{it}^s + \beta_{2i} \varepsilon_{it}^d \quad (3.11)$$

$$\Delta e_{it} = E_{t-1}(\Delta e_{it}) + \gamma_{1i} \varepsilon_{it}^s + \gamma_{2i} \varepsilon_{it}^d + \gamma_{3i} \varepsilon_{it}^e \quad (3.12)$$

Furthermore, short term interest rates are used to incorporate central bank policy in the system. Monetary policy may react to exchange rate fluctuations and then policy may affect exchange rates and domestic inflation. That way, the observed relationship between prices and exchange rates would take into account the central bank behavior rather than the direct influence of exchange rates on prices. As discussed by PARSLEY and POPPER (1998), taking into account monetary policy would improve significantly the estimation results of ERPT. Given this view, the last portion of the model consists of a central bank reaction function. The reaction function relates short-term nominal interest rates (r_{it}) to the previously cited variables in the model as central banks use the short-term rate as their monetary policy instrument.

$$\Delta r_{it} = E_{t-1}(\Delta r_{it}) + \lambda_{1i} \varepsilon_{it}^s + \lambda_{2i} \varepsilon_{it}^d + \lambda_{3i} \varepsilon_{it}^e + \lambda_{4i} \varepsilon_{it}^{mpi} + \lambda_{5i} \varepsilon_{it}^{ppi} + \lambda_{6i} \varepsilon_{it}^{cpi} + \varepsilon_{it}^r \quad (3.13)$$

Finally, the conditional expectations ($E_{t-1}(\cdot)$) in equations (3.7) to (3.13) is assumed to be replaced by linear projections on lags of the seven variables in the system. In a such framework, the model can be expressed and estimated as a VECM using a Cholesky decomposition to identify the shocks.¹⁷ As is well known, this identification technique can be sensitive to the ordering of variables. We have just seen in MCCARTHY

¹⁷Note that even though the data in this study have both cross-sectional and time-series aspects, the model will be estimated for each country separately. The differing institutions in each country are likely to lead to differences in the responses in each country.

(2007) model that the use of a recursive identification scheme implies that the identified shocks contemporaneously affect their corresponding variables and those variables that are ordered at a later stage, but have no impact on those that are ordered before. As a matter of fact, when the reduced-form residuals from the system do not display high cross correlations, the order of factorization makes little difference. From the variance-covariance matrix, the correlations between residuals are less than 0.3, with the notable exceptions of the exchange rate and import prices and between oil prices and producer prices. Nevertheless, given that we are aware of the possible sensitivity of the Cholesky approach to the ordering of the variables, we conduct a sensitivity analysis by computing the generalized impulse response functions, as introduced by PESARAN and SHIN (1998), where ordering of the variables does not matter (see Appendix C.6).

It is worth highlighting that our model differs slightly from that of MCCARTHY (2007). The author estimate first-difference VAR model ignoring the possibility of cointegration among the levels of the variables. However, if the levels of a time series are non-stationary, non-sense results may occur if the non-stationarity is ignored. Thus, we feel that it is more appropriate to retain the information contained in the levels of the variables and then derive impulse responses from the VECM which incorporate the long run relationships among the variables. Moreover, throughout the *single-equation* literature of ERPT, a proxy for foreign producers' costs was considered as a primary control variable. Along with this literature, we propose to include a measure of foreign costs as exogenous variables in our VECM. Doing so, we think that this gives more reliable estimates of pass-through.¹⁸ In the following sub-sections, we report responses of domestic prices to both exchange rate shocks (sub-section 6.2) and import price shocks (sub-section 6.3).

¹⁸One can think that oil prices should enter our VECM as an exogenous variables given the likely weakly exogeneity of this variable (see adjustments coefficients in Table 3.4). Since oil prices is considered as the most exogenous variable, it is sensible to order it first in the scheme (see HAHN (2003) and HÜFNER and SCHRÖDER (2002)).

6.2. Responses to exchange rate shocks

In this sub-section, we report the impulse responses of all stages of the distribution chain (import, producer and consumer prices) to exchange rate shocks. This gives us the opportunity to analyze how exchange rate fluctuations are propagated from one price stage to the next. Thus, first the responses of import prices to the different external shocks are discussed. Thereafter we turn to the responses of producer prices and finally to those of the consumer prices. Although the model is estimated in first differences, it is then transformed into levels so that cumulative price level responses are displayed over a time horizon of twelve quarters.¹⁹ All shocks are standardized to a 1% shocks to allow a comparison of the sensitivity to currency shocks across countries. The horizontal axis measures the time horizon in terms of quarters after the shock; the vertical axis measures the deviation in (log) prices from their baseline levels indicating the approximate percentage point change in the respective price index due to a 1% shock in the exchange rate (which corresponds to 1% depreciation), i.e. the percentage of the ERPT. In the second part of this sub-section, we discuss some macroeconomic determinants that may affect the degree of ERPT.

6.2.1. Impulse responses analysis

Figures 3.2 to 3.4 display respectively the responses of the import price, the producer prices and the consumer prices to a 1% exchange rate shock in each of the EA countries.²⁰ Also, in Table 3.6, we report the response of each price index at various horizons (after 0, 1, 4 and 8 quarters). As mentioned before, the robustness of the identification scheme adopted in our study is checked using generalized impulse response functions (PESARAN and SHIN (1998)). According to the response of consumer prices reported in Appendix C.6, our ERPT estimates are broadly robust, using *generalized* impulse responses, instead of the *orthogonalized* recursive ones, do

¹⁹This is the most relevant time period for our analysis and the effects thereafter in most cases are not significant.

²⁰Confidence intervals for the impulse response functions are estimated using the Bayesian Monte Carlo method employed by RATS with 1000 replications.

not change the broad pattern and magnitude of the transmission of exchange rate shocks to consumer prices.

Beginning with the impact of an exchange rate depreciation on import prices in Figure 3.2). As expected, the response is positive following one percent of currency depreciation with a considerable cross-country variation in our EA sample. The highest immediate response (namely 0 quarter in Table 3.6) is recorded in Italy roughly 0.62%, while the lowest is in Austria equal to 0.29%. Also, Italy has the fastest import prices reaction; there is a complete ERPT after a single quarter. Interestingly, we notice that, by the end of the first year, a complete pass-through was detected in 7 out of 12 EA countries. Comparing our results with previous studies, our estimates of pass-through seem to be higher. We think that differences in the results are owing to different econometric methods used to estimate pass-through. In a single-equation context, ANDERTON (2003) found that 0.50 to 0.70% of changes in the euro are passed-through to import prices (in the long-run) over 1989-2001. As is well known, contrary to the single-equation method, a CVAR model allow for system estimation where the endogenous variables are simultaneously determined. Simply ignoring such simultaneity, as is often done in single-equation approaches, would result in simultaneous equation bias. Furthermore, in a pricing chain model, CVAR model permits for underlying dynamic interrelations among prices at different stages of distribution and other variables which cannot be done within single-equation method. Thus, it is not surprising that import price pass-through in our CVAR analysis lies somewhat above those single-equation estimates. We pretend that CVAR models would provide more relevant measure of the extent of ERPT, since it gives us the opportunity to analyze how exchange rate fluctuations pass through the production process from the import of products to the consumer level. Besides, HAHN (2003) found that pass-through amounts to about 0.50% after three quarters for the whole euro area. However, the author estimated a VAR in first-differences which does not incorporate the long-run relationship. We think that the neglect of time-series properties of the data - non-stationarity and cointegration relationship - would explain the relative weakness of ERPT estimates in comparison to our study.

Figure 3.2: Response of import prices to 1% exchange rate shock

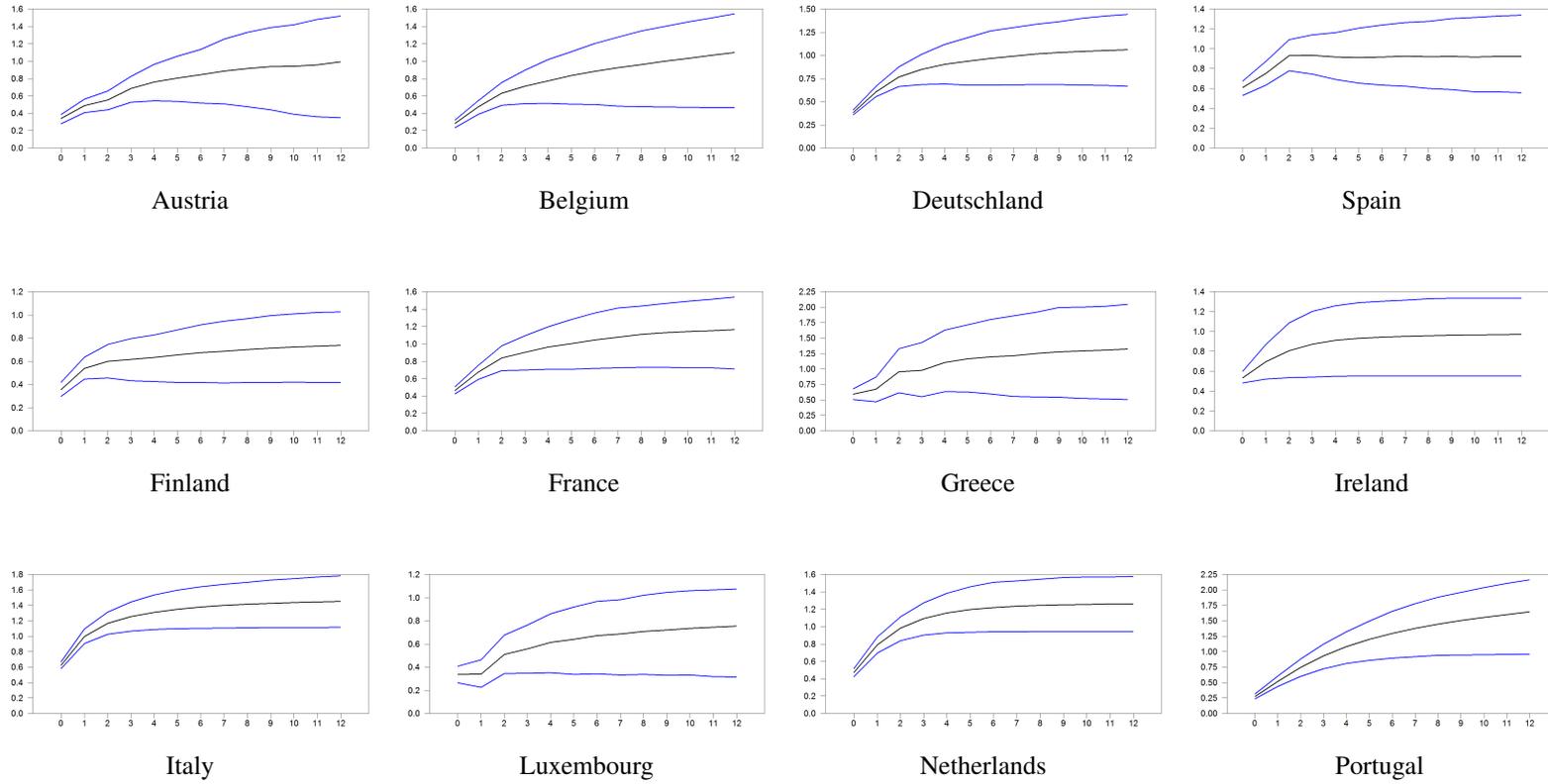


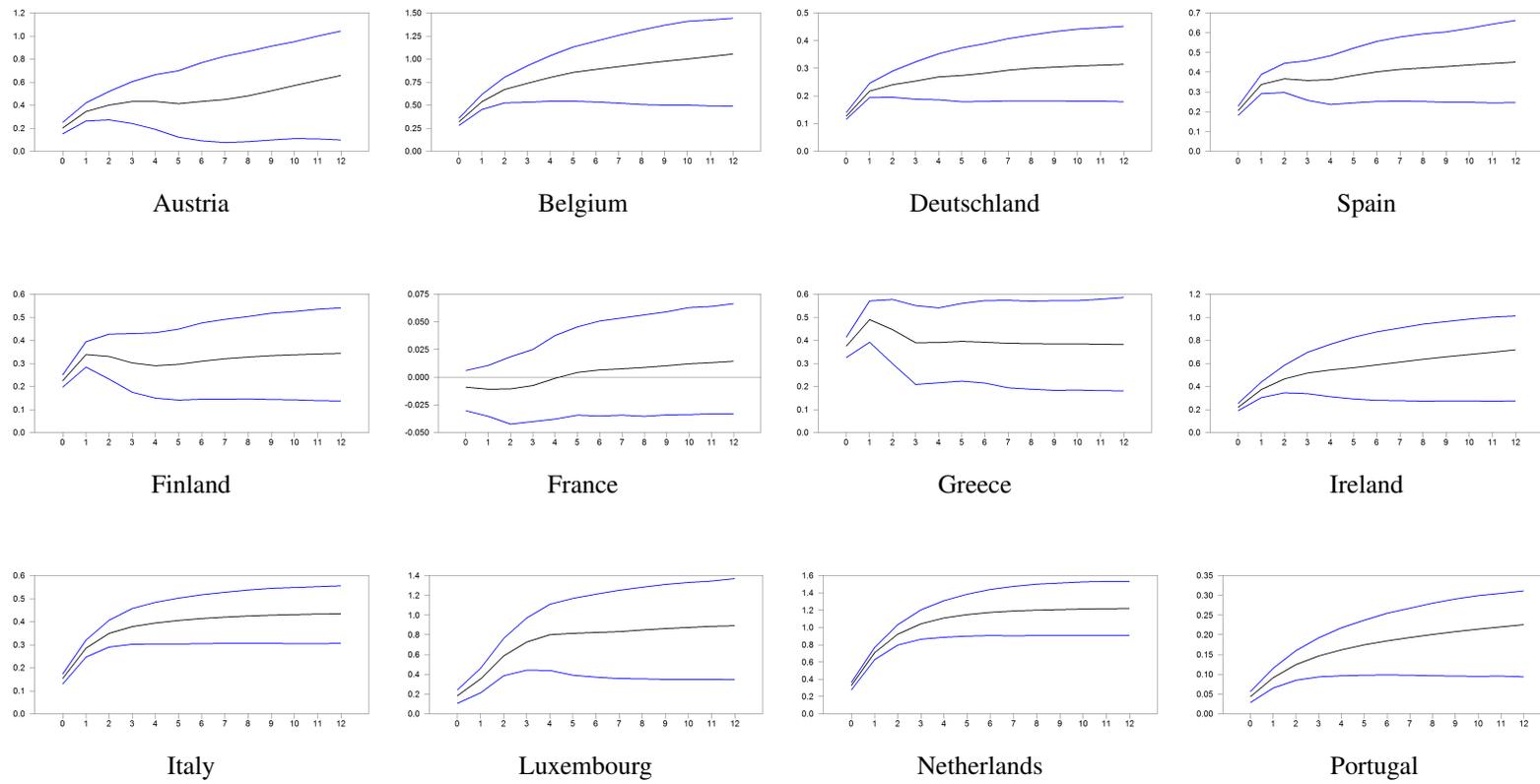
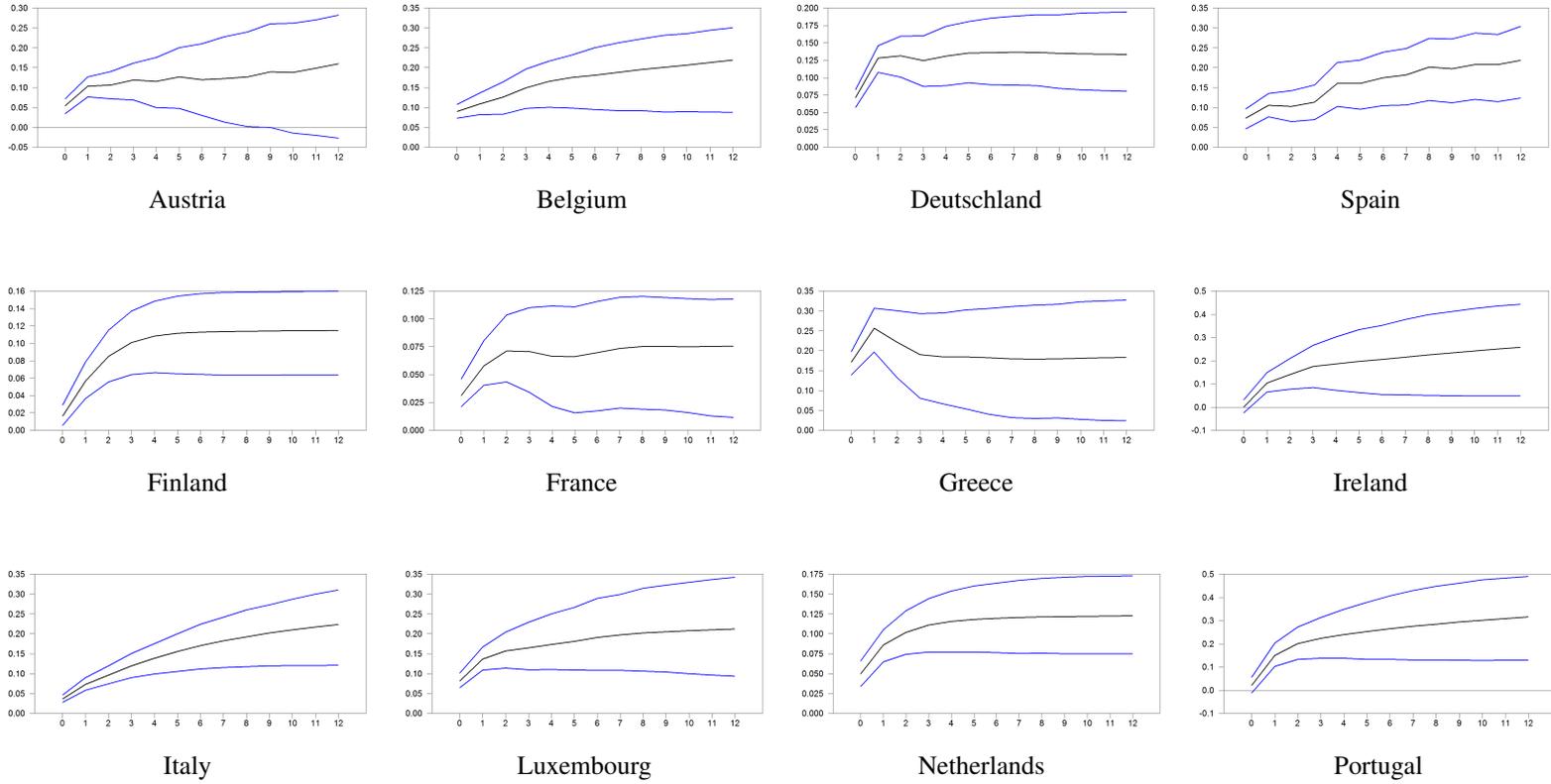
Figure 3.3: Response of producer prices to 1% exchange rate shock

Figure 3.4: Response of consumer prices to 1% exchange rate shock



As regards the response of producer prices, Figure 3.3 points out a more pronounced cross-country differences which is an expected phenomenon. We find that ERPT is surprisingly not significant (slightly negative) in France, very weak in Portugal (not exceeding 0.2% within two years), and complete in Netherlands within only one year. The higher responsiveness of producer prices in Netherlands was confirmed by MCCARTHY (2007). The authors found that pass-through to be particularly large in Belgium and the Netherlands in comparison with the rest of his sample of nine industrialized countries. For the euro area, HAHN (2003) report that 1% exchange rate shock is passed-through on producer prices by 0.10% percent after one quarter, by 0.28% after one year, and amounts to about 0.30% percent after three years. These estimates are close to those found for Germany in our study.

Finally, we focus on the pass-through of a 1% depreciation of exchange rate to consumer prices. Our results reveal a weak response in most of the EA countries with as usual a wide dispersion of rates of pass-through. The highest immediate effect can be observed in Greece with a consumer price index increase of 0.12%, which increases to 0.20% after one year. While the lowest estimated pass-through is found in France, the response of consumer prices does not exceed 0.08% across the different time horizons. For France, this result is not surprising since the response in the previous stage of distribution, i.e. producer price, was not significant. Moreover, we can say that results from impulse response functions corroborate what we find in the cointegration analysis. According to this latter, the highest long-run ERPT was found in Greece and Portugal, while the lowest was recorded in France and Finland. In fact, the weakness of consumer prices responsiveness was confirmed throughout VAR literature (see summary of VAR literature in Table 3.1). In their study, HÜFNER and SCHRÖDER (2002) found that the highest effect is observed in the Netherlands which is equal to 0.12% within one year. Also, for France, HÜFNER and SCHRÖDER (2002) report ERPT estimates as low as those found in our study. The pass-through of one percent depreciation on the consumer prices in the first year is roughly 0.07%. It should be note that our study is close to HÜFNER and SCHRÖDER (2002) who use a VECM approach incorporating long-run relationships between variables. The main difference vis-à-vis our study is that we include a measure of foreign costs as exogenous variable which is considered as a primary variables throughout ERPT literature.

Table 3.6: Impulse response along the distribution chain of pricing

Accumulated response of import prices to 1% exchange rate shock						
Response horizon	Austria	Belgium	Germany	Spain	Finland	France
0	0,289	0,297	0,359	0,653	0,411	0,481
1	0,442	0,486	0,614	0,819	0,591	0,697
4	0,710	0,775	0,871	0,981	0,654	0,946
8	0,835	0,865	0,955	0,959	0,694	1,033
Response horizon	Greece	Ireland	Italy	Luxembourg	Netherlands	Portugal
0	0,508	0,555	0,615	0,342	0,478	0,360
1	0,587	0,732	1,016	0,344	0,810	0,616
4	1,015	0,913	1,297	0,570	1,172	1,051
8	1,067	0,927	1,373	0,633	1,229	1,216

Accumulated response of producer prices to 1% exchange rate shock						
Response horizon	Austria	Belgium	Germany	Spain	Finland	France
0	0,197	0,318	0,119	0,244	0,229	-0,005
1	0,340	0,549	0,205	0,390	0,346	-0,007
4	0,405	0,795	0,253	0,425	0,290	-0,001
8	0,401	0,867	0,270	0,486	0,315	0,008
Response horizon	Greece	Ireland	Italy	Luxembourg	Netherlands	Portugal
0	0,459	0,270	0,176	0,155	0,279	0,052
1	0,646	0,432	0,309	0,318	0,654	0,105
4	0,518	0,571	0,418	0,721	1,048	0,171
8	0,421	0,570	0,440	0,728	1,113	0,194

Accumulated response of consumer prices to 1% exchange rate shock						
Response horizon	Austria	Belgium	Germany	Spain	Finland	France
0	0,056	0,105	0,062	0,080	0,014	0,046
1	0,103	0,126	0,125	0,111	0,054	0,079
4	0,113	0,179	0,123	0,164	0,103	0,075
8	0,106	0,192	0,134	0,199	0,105	0,077
Response horizon	Greece	Ireland	Italy	Luxembourg	Netherlands	Portugal
0	0,180	0,009	0,049	0,087	0,073	0,058
1	0,263	0,107	0,090	0,141	0,111	0,184
4	0,203	0,182	0,152	0,172	0,140	0,251
8	0,191	0,200	0,195	0,177	0,143	0,278

Note: Response horizon 0, 1, 4 and 8 correspond, respectively, to immediate, one quarter, one year and two years response after the initial shock.

For the non-US G7 countries, CHOUDHRI, FARUQUEE, and HAKURA (2005) estimate a first-difference VAR model that contained exogenously determined foreign variables, namely foreign consumer price index and foreign interest rate. In a very similar framework without including foreign exogenous variables, FARUQUEE (2006)

reports a quite weak pass-through to consumer prices (see Table 3.1). Thus, we pretend that is a sensible way to enter foreign exogenous variables - such as foreign prices or costs - when estimating the extent of pass-through within VAR framework.

Otherwise, the response of the consumer prices to the exchange rate shock is found to be weaker than that of the producer prices. Imports as intermediate goods that need to go through production or distribution processes before they are consumed by households. The production or distribution channels can dampen the effect of exchange rate changes and account for a low pass-through to consumer prices. Also, our results point out that pass-through declines along the distribution chain with the largest effect occurring in import prices. This decline is due to a smaller fraction of goods affected by external factors in the price indices at later stages of the distribution chain. In other words, the fraction of goods that are affected by exchange rate shocks seems to decrease along the distribution chain, pointing to a declining pass-through. For example, the share of tradables, that are likely to be more prone to external shocks than non-tradables (services), tends to decrease in price indices along the distribution chain. Furthermore, assuming that shocks are, at least partially, passed-through via previous stages, thus, accumulation over different stages basically implies a decline in the pass-through along the distribution chain. Another line of argumentation used in pass-through literature to explain the observed smaller pass-through to consumer prices compared to import prices: the presence local distribution costs, the extent of imported inputs being used for domestic production (see BURSTEIN, EICHENBAUM, and REBELO (2005)) or the optimal pricing strategies of foreign producers and domestic wholesalers/retailers (BACCHETTA and VAN WINCOOP (2002)).

6.2.2. *Factors influencing ERPT*

In order to explain the cross-country differences detected from impulse responses along distribution chain, we introduce some macroeconomic determinants which can explain the differences in pass-through estimates in our 12 EA countries. To this end, we examine the Spearman rank correlation statistic between the impulse responses for different prices (import prices, producer prices and consumer prices) at various horizons and some factors expected to influence pass-through. There are various theoretical

arguments have been made for cross-country differences in exchange rate pass-through rates. In our study we analyze the differences in the degree of pass-through into import prices across the five following determinants: (1) mean of import share or degree of openness (imports as a percentage of domestic demand) over the sample period 1980-2010; (2) exchange rate persistence measured as the impulse response at the 8-quarter horizon of the exchange rate to its own standardized shock;²¹ (3) exchange rate volatility measured by the standard deviation of quarterly percentage changes in the exchange rate ($\sigma_{\Delta e}$);²² (4) inflation level as the mean of the year-on-year quarterly inflation rate over sample period; (5) inflation volatility as the standard deviation of the year-on-year quarterly inflation rate over sample period. Results of rank correlation are reported in Table 3.7.

Table 3.7: Rank correlation between ERPT and Selected Variables

Macroeconomic Determinants	Response horizon			
	0	1	4	8
(a) Impulse response of import prices				
<i>Import Share</i>	-0,286	-0,153	-0,118	-0,132
<i>Exchange rate persistence</i>	0,629**	0,769***	0,657**	0,643**
<i>Exchange rate volatility</i>	-0,655**	-0,566*	-0,325	-0,398
<i>Inflation</i>	0,622**	0,3147	0,601**	0,538*
<i>Inflation volatility</i>	0,594**	0,3287	0,531*	0,538*
(b) Impulse response of PPI				
<i>Import Share</i>	0,587**	0,748***	0,769***	0,748***
<i>Exchange rate persistence</i>	-0,384	-0,370	-0,244	-0,153
<i>Exchange rate volatility</i>	-0,655**	-0,566*	-0,384	-0,398
<i>Inflation</i>	0,511*	0,517*	0,2081	0,1617
<i>Inflation volatility</i>	0,1307	-0,1818	-0,2657	-0,3077
(c) Impulse response of CPI				
<i>Import Share</i>	0,132	0,062	0,335	0,475
<i>Exchange rate persistence</i>	-0,138	0,161	0,554*	0,676**
<i>Exchange rate volatility</i>	-0,655**	-0,566*	-0,384	-0,398
<i>Inflation</i>	0,675**	0,861***	0,506*	0,392
<i>Inflation volatility</i>	0,269	0,613**	0,581**	0,527*

Note: *, ** and *** denote significance level at 10%, 5% and 1% respectively.

²¹As defined by MCCARTHY (2007).

²²We adopt the same exchange rate volatility proxy employed by BARHOUMI (2006).

Concerning imports prices, as expected, the extent of pass-through is positively correlated with the persistence of exchange rate, inflation level and inflation volatility with a significant relationship. This latter result is in line with TAYLOR (2000) who has put forward the hypothesis that the responsiveness of prices to exchange rate fluctuations depends positively on inflation environment. Also, TAYLOR (2000) explained that a higher perceived persistence of exchange rate shocks would entail a larger extent of pass-through. However, Table 3.7 reports a significant negative correlation between imports prices response and exchange rate volatility. In fact, ERPT literature is not conclusive with respect to the relationship between the volatility of exchange and the degree of pass-through. On one hand, there is a strand of literature supporting the presence of a negative correlation. Greater exchange rate volatility may make importers more wary of changing prices and more willing to adjust profit margins, thus reducing measured pass-through (see MANN (1986)). This hypothesis was confirmed by some empirical studies (see WEBBER (1999) and BARHOUMI (2006) among others). On the other hand, it is expected that import prices responsiveness would be higher when volatility of exchange rate is larger. As pointed by DEVEREUX and ENGEL (2002), the relative stability of importing country's currency plays a substantial role in determining pass-through. Countries with low relative exchange rate variability would have their currencies chosen for transaction invoicing (LCP strategy). CAMPA and GOLDBERG (2005) found that exchange rate volatility affects in a statistically significant way the degree of pass-through. We see that our results are rather in line with the first hypothesis. As for import share, the relationship with ERPT is very weak with a wrong negative sign. This is not surprising since the greater openness of a country may be an indicative of increased foreign competitive pressures limiting exchange rate transmission. Empirically, this was confirmed by CA'ZORZI, KAHN, and SÁNCHEZ (2007) and MCCARTHY (2007) using first-difference VAR model.

As regards producer prices, the results are quite different in comparison to import prices. We point out a positive relationship with degree of openness which is statistically significant throughout different time horizons. Imported goods as intermediate goods have to go through production or distribution processes before they reach consumers. Thus, higher import shares could be correlated with a greater producer price response. Inflation environment have the expected correlation and it is statistically significant in a

shorter horizons. For exchange rate volatility, the relationship is rather negative as in the case of import prices. However, exchange rate persistence and inflation volatility display no strong correlation with the producer price response. Finally, regarding consumer prices response, the results are quite similar to those for import prices. The exception is the degree of openness which is found to be positively correlated with the ERPT to consumer prices although the relationship is not statistically significant. These results are consistent with the empirical literature dealing with the so-called “second-stage pass-through”. In a panel of 71 countries, CHOUDHRI and HAKURA (2006) show that ERPT is positively correlated to the average of inflation rate and the inflation and exchange rate volatility, but no significant role for the degree of openness was founded.

In the end of this sub-section, we want to compare results derived from impulse responses with those obtained from the earlier cointegration analysis (see section 5 above). Summary of estimates of ERPT to consumer price resulting from the two analyses are displayed in Table 3.8. The first curious result is that estimates from the impulse response function analysis are somewhat lower compared to cointegration analysis. This is not surprising since cointegration analysis provide a longer time horizon. Our earlier results in section 5 revealed low loading factors indicating a slow adjustment of consumer prices to their long-run equilibrium. Thereby, the adjustment process is not fully completed during the considered time horizon in the impulse response analysis (8 quarters), and it is expected that the long-run effects found in the cointegration analysis should effectively be somewhat higher. Nevertheless, we note that impulse response estimates at 8 quarters are extremely close to the cointegration estimates for Finland and Germany. As regards the magnitude of the pass-through, we can say that results from impulse response functions corroborate in some extent what we find in the cointegration analysis. The highest ERPT estimates were found in Greece and Portugal, while the lowest was recorded in France and Finland. Finally, concerning factors influencing the rate of pass-through, we point out the important role of inflation level, inflation volatility and (in a lesser extent) the exchange rate persistence in explaining the cross-country differences in the long-run. Higher inflation level or volatility and more exchange rate persistence are correlated with a greater response of consumer prices in the long-run.

To sum up, our results show a higher pass-through to import prices with a complete pass-through detected in roughly half EA countries after one year. These results are relatively large compared to single-equation literature. The magnitude of the pass-through of exchange rate shocks decline along the distribution chain of pricing with the modest effect is recorded with consumer prices. Also, referring to the magnitude of the pass-through, we can say that results from impulse response functions corroborate in some extent what we find in the cointegration analysis. The highest ERPT estimates were found in Greece and Portugal, while the lowest was recorded in France and Finland. When assessing possible reasons for cross-country differences in the ERPT, inflation level, inflation volatility and exchange rate persistence are the main macroeconomic factors that influencing the degree of pass-through almost along the distribution pricing chain. The exchange rate volatility is surprisingly negatively correlated with response of different prices index.

Table 3.8: Summary of ERPT to consumer price index (CPI)

Country	Response of CPI to 1% exchange rate shock			Cointegration
	1 quarter	4 quarters	8 quarters	Long-run
Austria	0,103	0,113	0,106	0,248
Belgium	0,126	0,179	0,192	0,282
Germany	0,125	0,123	0,134	0,169
Spain	0,111	0,164	0,199	0,337
Finland	0,054	0,103	0,105	0,117
France	0,079	0,075	0,077	0,166
Greece	0,263	0,203	0,191	0,576
Ireland	0,107	0,182	0,200	0,397
Italy	0,090	0,152	0,195	0,352
Luxembourg	0,141	0,172	0,177	0,339
Netherlands	0,111	0,140	0,143	0,298
Portugal	0,184	0,251	0,278	0,833
Spearman rank correlation				
<i>Import Share</i>	0,063	0,336	0,476	-0,140
<i>Exchange rate persistence</i>	0,161	0,554*	0,676**	0,469
<i>Exchange rate volatility</i>	-0,566*	-0,385	-0,399	-0,070
<i>Inflation</i>	0,8615***	0,506*	0,393	0,643**
<i>Inflation volatility</i>	0,613**	0,581**	0,527*	0,664**

Note: *, ** and *** denote significance level at 10%, 5% and 1% respectively.

6.3. Responses to import price shocks

In this sub-section, we focus on the responses of domestic prices; i.e. producer prices and consumer prices, to 1% shock in import prices. This analysis is of a great interest since it provides insights how shocks are propagated from one price stage to the next. We have seen in the identification scheme that the import price shock is estimated given past values of all the variables plus the current value of oil prices, the real GDP, and the exchange rate. Results of pass-through of import prices to domestic prices are reported in Figure 3.5 and 3.6, and in Table 3.9.

Beginning with producer prices, as expected the pass-through is positive in most of EA countries but not significant for some countries, namely Spain, France, Greece and Ireland. The highest response are identified in Belgium and Netherlands, this may explain why ERPT to producer prices is found to be higher in these EA countries. Especially, for Belgium, the pass-through of 1% increase in import prices raise producer prices more than 1% within one year. Similarly, MCCARTHY (2007) reports that response are particularly large in Belgium, with the pass-through eventually exceeding 1%. For the whole EA, HAHN (2003) found that the impact of a one percent increase in non-oil import prices on producer prices is extremely large. In the first quarter the pass-through amounts to 0.22%, increasing to 0.61% after one year. Also, we note that the effect of import prices are broadly weak compared to exchange rate shocks (see previous section 6.2) with the exception of Germany. Nevertheless, HAHN (2003) points out that pass-through of oil prices and Exchange rate to producer prices are smaller than impact of import prices. According to the author, this may be due to a higher perceived persistence of the import price shocks. While exchange rate and oil price shocks are known to be pretty volatile, import price shocks are likely to contain the more persistent external sources of variation. Besides, according to our results, the response of producer prices has insignificantly the wrong (negative) sign in France. These negative coefficients consolidate what we found in the previous section, namely the insignificant (negative) ERPT to producer prices. In fact, when domestic currency is depreciating, domestic producer and wholesalers may stop stocking foreign products (as intermediate goods) since their price becomes too high. Thus, a substitution effect occurs and producer prices will be more insulated from imports prices changes. This

may explain why the response of producer prices is not significant in some EA countries such as France.

The response of consumer prices to import price shocks is also positive for most EA countries but statistically significant only for the half of our sample (with negative sign for Spain and Ireland). This outcome would explain the weakness of ERPT to consumer prices in our sample. The highest effect is detected in Greece and Portugal which is a natural result as these countries has the highest degree of pass-through of exchange rate. This latter result may be considered as an evidence of weak *pricing-to-market* behavior in the domestic markets of Portugal and Greece in comparison to the rest of EA members. Furthermore, we note that for 8 out of 12 EA countries the effect of import prices on consumer prices are smaller than exchange shocks.

Table 3.9: Impulse response along the distribution chain of pricing

Accumulated response of producer prices to 1% increase in import prices						
Response horizon	Austria	Belgium	Germany	Spain	Finland	France
0	0,160	0,342	0,042	0,114	0,159	-0,092
1	0,174	0,671	0,043	0,072	0,221	-0,042
4	0,341	1,157	0,402	0,088	0,224	-0,093
8	0,229	1,287	0,486	0,090	0,225	-0,091
Response horizon	Greece	Ireland	Italy	Luxembourg	Netherlands	Portugal
0	0,176	0,227	0,124	0,232	0,114	0,085
1	0,166	0,257	0,172	0,437	0,378	0,118
4	0,153	0,335	0,130	0,639	0,819	0,184
8	0,476	0,305	0,120	0,500	0,893	0,223
Accumulated response of consumer prices to 1% increase in import prices						
Response horizon	Austria	Belgium	Germany	Spain	Finland	France
0	-0,031	0,070	0,073	0,014	-0,001	0,134
1	0,031	0,174	0,017	-0,090	0,012	0,191
4	0,024	0,353	0,120	-0,070	0,006	0,217
8	0,034	0,373	0,121	-0,077	0,001	0,239
Response horizon	Greece	Ireland	Italy	Luxembourg	Netherlands	Portugal
0	0,263	-0,100	0,031	0,013	0,013	0,206
1	0,379	-0,150	0,064	0,050	0,047	0,250
4	0,540	-0,232	0,047	0,079	0,065	0,431
8	0,631	-0,256	0,048	0,045	0,066	0,514

Note: Response horizon 0, 1, 4 and 8 correspond, respectively, to immediate, one quarter, one year and two years response after the initial shock.

Figure 3.5: Response of producer prices to 1% increase in import prices

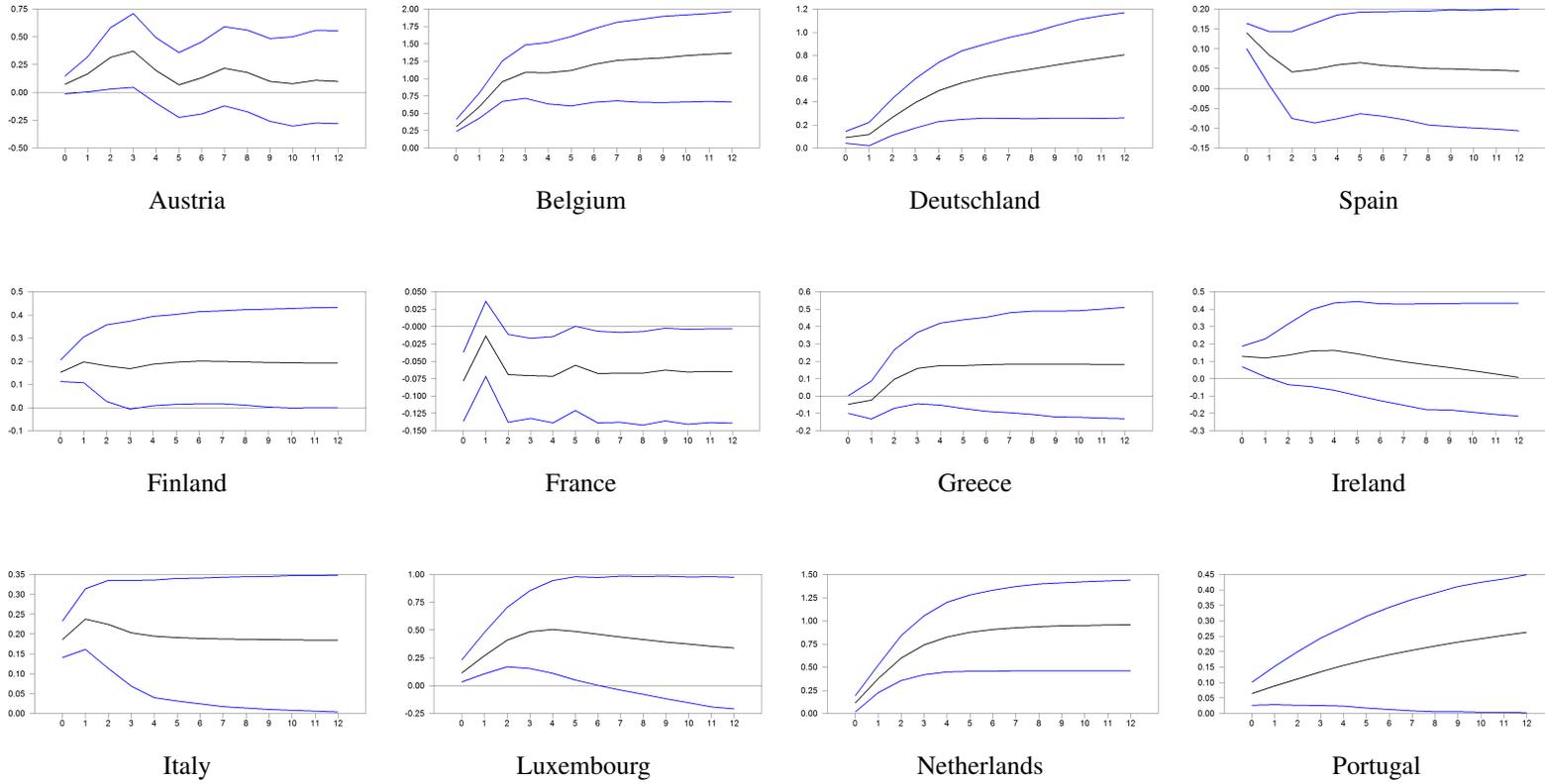
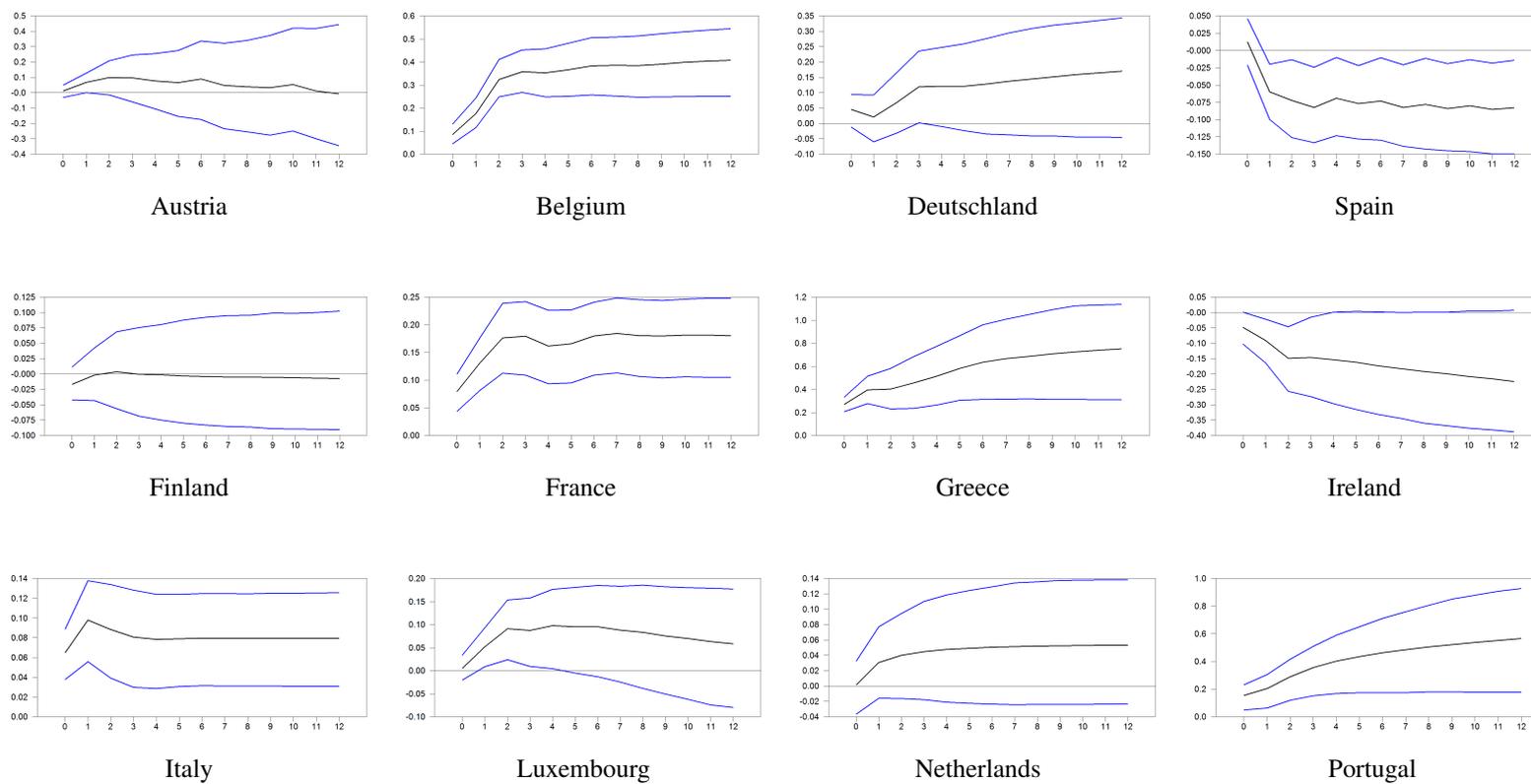


Figure 3.6: Response of consumer prices to 1% increase in import prices

As is well-known, exchange rate changes may be transmitted *directly* to prices consumer through the price of imports. In the case of depreciation, domestic-currency price of the imported good will rise in proportion. This change in import prices is then likely to translate into changes in the producer and consumer prices if producers raise their prices in line with the increase in import prices. On the other hand, currency depreciation may affect *indirectly* consumer prices through changes in the composition of demand or in the levels of aggregate demand and wages. A depreciation of the exchange rate makes domestic products relatively cheaper for foreign buyers, and as a consequence exports and aggregate demand will rise and induce an increase in the domestic price level. At the same time, the increase of domestic demand also leads to a higher demand for labour and, potentially, to rising wages, which will in turn be reflected in higher prices. Consequently, it is expected that exchange shocks would have a higher effect than import prices shocks as showed by our results.

6.4. Variance decompositions

It is known that impulse responses trace the effects of a shock to one endogenous variable on to the other variables in the VECM system, allowing us to estimates of the effect of exchange rate and import price shocks on domestic producer and consumer prices. However, impulse responses do not enable us to determine the importance of these “external” shocks for domestic price fluctuations over the sample period. To get additional insights on this, we examine the variance decompositions of the different price indices, i.e. import prices, producer prices and consumer prices. Variance decompositions separate the variation in an endogenous variable into the component shocks to the VECM system. Thus, the variance decomposition provides information about the relative importance of each random innovation in affecting the variables in the system. In other words, variance decompositions indicate the percentage contribution of the different shocks to the variance of the h -step ahead forecast errors of the variables. Hence, the relative importance of the different external shocks for the development of the price indices could be assessed.

Tables 3.10 to 3.12 summarizes the results on the variance decompositions of import, producer and consumer prices over a forecast horizon of 0, 1, 4, and 8

quarters. For import prices, we report only the contribution exchange rate shocks, while for producer and consumer prices, the contribution of “external shocks”, i.e. of exchange rate shocks, and import price shocks, are displayed. Also, in the lower part of different Tables, we set out the rank correlations between the percentage of one price index variance attributed to exchange rate shocks and the different macroeconomic determinants listed in the sub-section 6.2.

Table 3.10: Percentage of import price forecast variance attributed to exchange rate shocks

Country	Forecast horizon			
	0	1	4	8
Austria	1,84	9,80	10,75	10,61
Belgium	13,26	10,10	11,19	11,74
Germany	16,09	16,51	16,52	16,53
Spain	26,97	19,60	19,23	19,18
Finland	14,98	14,03	12,97	12,88
France	25,37	27,76	31,72	32,26
Greece	65,22	59,03	52,83	52,73
Ireland	60,02	55,40	54,85	54,85
Italy	41,79	43,95	43,30	43,12
Luxembourg	0,00	0,94	5,01	4,99
Netherlands	21,64	24,54	25,80	25,79
Portugal	39,65	42,17	46,75	47,38
Spearman rank correlation coefficient with:				
<i>Import Share</i>	-0,6364**	-0,6294**	-0,5734*	-0,5755*
<i>Exchange rate persistence</i>	0,5874*	0,5455*	0,5874**	0,5874**
<i>Exchange rate volatility</i>	-0,439	-0,474	-0,460	-0,461
<i>Inflation</i>	0,6713**	0,5874*	0,5734*	0,584*
<i>Inflation volatility</i>	0,8042***	0,7762***	0,7902***	0,785***

Note: *, ** and *** denote significance level at 10%, 5% and 1% respectively..

Beginning by examining the variance decomposition of import price (Tables 3.10). Again, results differ across countries. It can be seen that exchange rate shocks explain a fairly large part of the fluctuation of import prices especially in Greece, Ireland, Italy and Portugal. In these countries, the shares range from over 40% to 60%. While for other countries, like Austria, Belgium, Finland and Luxembourg, the importance of exchange rate does not exceed 15%. We point out that the percentage of contribution exchange rate shocks increases for most countries as the forecast horizon increases, as a proof of gradual adjustment of import prices; it takes time until changes in exchange rate are reflected in the import prices. Thereafter, we look to the correlation between the

percentage of import price variance attributed to exchange rate and factors influencing ERPT. Our results reveal the same conclusion when using impulse responses: the contribution of exchange rate to import prices fluctuations is negatively correlated to the degree of openness and exchange rate volatility (albeit not significant), while the relationship is strongly significant (with a positive sign) with exchange rate persistence, inflation level and inflation volatility.

Table 3.11: Percentage of producer prices forecast variance attributed to external shocks

Country	Forecast horizon			
	0	1	4	8
Austria	1,84	9,80	10,75	10,61
Belgium	13,26	10,10	11,19	11,74
Germany	16,09	16,51	16,52	16,53
Spain	26,97	19,60	19,23	19,18
Finland	14,98	14,03	12,97	12,88
France	25,37	27,76	31,72	32,26
Greece	65,22	59,03	52,83	52,73
Ireland	60,02	55,40	54,85	54,85
Italy	41,79	43,95	43,30	43,12
Luxembourg	0,00	0,94	5,01	4,99
Netherlands	21,64	24,54	25,80	25,79
Portugal	39,65	42,17	46,75	47,38
Spearman rank correlation coefficient with:				
<i>Import Share</i>	-0,6364**	-0,6294**	-0,5734*	-0,5755*
<i>Exchange rate persistence</i>	0,5874*	0,5455*	0,5874**	0,5874**
<i>Exchange rate volatility</i>	-0,439	-0,474	-0,460	-0,461
<i>Inflation</i>	0,6713**	0,5874*	0,5734*	0,584*
<i>Inflation volatility</i>	0,8042***	0,7762***	0,7902***	0,785***

Note: External shocks denote exchange rate and import shocks together. *, ** and *** denote significance level at 10%, 5% and 1% respectively..

With regard to the variance of producer prices, the contribution of external factors - exchange rates and import prices - is still high for the same group of countries, namely Greece, Ireland, Italy and Portugal (see Table 3.11). It should be noted that these countries have the highest long-run ERPT according to the cointegration analysis. Results reveals that external factors explain from 40% to 65% of producer prices forecast variance in the mentioned countries, which is a quite high contribution compared to the other shocks that may heating the economy (such as supply or demand shocks). The contribution in the other countries is more modest, especially for Austria and Luxembourg. For the whole euro area, HAHN (2003) found that between 5% to 20%

of the variance are accounted for by exchange rate and import price shocks respectively. This result masks the wide dispersion between EA countries in terms of the importance of exchange rate and import prices shocks. Besides, we find that the percentage of producer prices variance attributed to external factors tend to be higher for countries with higher exchange rate persistence, inflation level and inflation volatility. The relationship with import share is still having the wrong negative sign.

Finally, we focus on the importance of external shocks for consumer prices fluctuations. In contrast to producer prices, the influence of external factors on consumer prices variance is weak. In most of EA countries, exchange rate and import prices shocks explain less than 18% of the variance of the consumer prices. This percentage tends to increase as the forecast horizon increases since it takes time until changes in the external factors are reflected in the consumer prices. Again, Greece and Portugal have the largest contribution of external shocks to consumer prices fluctuations. This may explain once again why ERPT to consumer prices are higher in the two countries compared to the rest of EA members. Otherwise, as usual, the differences across EA countries appear to be positively related with inflation level, inflation volatility and exchange rate persistence throughout time horizons. For the degree of openness the relationship is as usual negative. In fact, it is expected that the more country is open, the more exchange rate changes affect domestic prices. In a more open economy, with larger presence of imports and exports, a given depreciation would have a larger effect on prices. Thus, the most immediate connection between the two variables is positive. However, ROMER (1993) provided a theoretical explanation why inflation could be negatively correlated with openness, showing how openness puts a check on inflationary pressure. In this sense, inflation could be negatively correlated with openness. As explained by CA'ZORZI, KAHN, and SÁNCHEZ (2007), the existence of two mechanisms going in opposite directions may lead to a puzzling result and the overall sign of the correlation between pass-through and openness can thus be either positive or negative.

In summary, the variance decompositions indicate that external factors explain only a modest proportion of the forecast variance of domestic consumer prices over 1980-2010, while this contribution is to some extent high in Portugal and Greece. For the latter countries, this would explain why ERPT to consumer is relatively large compared to the other EA members. Mainly, three macroeconomic factors - inflation level, inflation

volatility and exchange rate persistence - are found to be crucial in explaining the cross-country differences regarding the influence of external shocks. The degree of openness is surprisingly negatively correlated with the contribution of external shocks.

Table 3.12: Percentage of consumer prices forecast variance attributed to exchange rate and import price shocks

Country	Forecast horizon			
	0	1	4	8
Austria	5,42	5,44	8,61	8,93
Belgium	4,08	3,86	3,87	3,87
Germany	3,60	4,23	6,56	6,58
Spain	12,11	10,56	10,36	10,24
Finland	2,12	2,81	4,19	4,17
France	0,70	2,75	9,80	10,04
Greece	9,52	10,60	15,98	16,51
Ireland	2,84	3,37	3,55	3,55
Italy	8,76	8,40	8,77	8,77
Luxembourg	5,88	6,73	9,56	9,44
Netherlands	0,22	3,79	4,21	4,22
Portugal	15,14	16,53	16,93	17,06
Spearman rank correlation coefficient with:				
<i>Import Share</i>	-0,3427	-0,3357	-0,6154**	-0,5734*
<i>Exchange rate persistence</i>	0,6464**	0,7050**	0,4996*	0,4796
<i>Exchange rate volatility</i>	0,1338	0,2017	0,3979	0,4124
<i>Inflation</i>	0,6503**	0,5385*	0,5734*	0,5315*
<i>Inflation volatility</i>	0,6475**	0,7102***	0,7090***	0,7104***

Note: External shocks denote exchange rate and import shocks together. *, ** and *** denote significance level at 10%, 5% and 1% respectively.

6.5. Testing for the recent decline in ERPT

In this sub-section, we investigate, whether the ERPT to consumer prices has changed over our sample period 1980-2010. Empirical literature has put forth the decline of rates of pass-through in major of industrialized countries (see *inter alia* GAGNON and IHRIG (2004)). Given the different developments experienced by the EA members (institutional arrangements (such as the introduction of the single currency in 1999), convergence of inflation rates, monetary and financial shocks (such as 1992/1993 ERM crises)), we examine the possible existence of structural shift in response of consumer prices to exchange rate shocks.

When assessing the stability of ERPT to consumer prices, we can speculate that the introduction of the euro, as a major economic event, would entail a changing in the behavior the exchange rate transmission. The literature raised a number of reasons why the rate of pass-through may have changed for the EA members as a result of entering the monetary union. Namely, the introduction of the single European currency has changed the competitive conditions by decreasing the share of trade exposed to exchange rate fluctuations. Also, the advent of the euro as well established currency in the 2000's, creating a single market for exporters, has favoured an expansion of the euro as a currency of denomination of its external trade. Referring to these factors, one can think that ERPT has declined in monetary union members following that date. As matter of act, empirical literature does not provide a strong evidence of structural break in pass-through coefficients since the creation of the euro area. In a set of studies, *CAMPA and GOLDBERG (2005, 2002)*, *CAMPA, GOLDBERG, and GONZÁLEZ-MÍNGUEZ (2005)* and *CAMPA and GONZÁLEZ (2006)* have tested the presence of structural break in the vicinity of the introduction of the common currency. Their results do not support the view that ERPT has declined around the date of the creation of the euro.

We have seen in the previous section that inflation environment (inflation level and inflation volatility) is an important macroeconomic factor influencing the ERPT. As argued by *TAYLOR (2000)*, the transition to the low inflation environment in many industrialized countries has successfully reduced the degree of pass-through to domestic prices. For the EA countries, the inflation convergence process has started before the adoption of the single currency, and more exactly, after the implementation of the Maastricht treaty.²³ Since higher inflation levels and volatility contribute to higher degree of pass through, countries that have experienced reduction in inflation and nominal volatility may have seen a significant lowering in pass through elasticities. Thus, for EA countries, we assume that a break exists and it would take place in the vicinity of the first stage of the EMU (in July 1990). To address this issue, we perform Chow test for structural change designed for multiple time series, as introduced by *CANDELON and LUTKEPOHL (2001)*, assuming an exogenously imposed break point

²³Among the Maastricht criteria for joining the EMU, each country's inflation in 1997 had to be less than 1.5 percentage points above the average rate of the three European countries with the lowest inflation over the previous year.

around the third quarter of 1990.²⁴ According test results reported in Table C.5 in Appendix C.5, there is a strong evidence of structural break around the starting of the first stage of the EMU for all EA countries.²⁵

To provide further insights on the changing behavior of ERPT, we use a simple strategy of reestimating our CVAR pricing chain model over a shorter sample period that does not include the 1980s, i.e. between 1990-2010. After deriving the impulse response of consumer prices to exchange rate shocks, this allows us to check the differences between the responses estimated over the whole sample (1980-2010) and those estimated over the shorter sub-sample (1990-2010) as in Figure 3.7. Almost all of EA countries (with few exceptions) show that exchange rate seems to have a less inflationary effect during last twenty years. According to impulse responses, there is an evidence of a general decline in rates of pass-through in most of euro zone countries. These findings confirm the presence of structural break as shown by chow tests. Given that inflation environment is an important determinant of ERPT, it is an expected result that the decline in response of consumer prices coincided with the steady reduction of inflation rates during the 1990's. This result is more apparent for the "peripheral" EA countries, namely Greece, Ireland, Portugal and Spain.²⁶

6.6. Historical Decompositions

In This sub-section, we use historical decompositions to examine the role played by the external shocks in the development of the consumer prices during two sub-sample periods: during the first and second stage of EMU (1990:1-1998:4); and since the creation of the euro till the end of our time sample (1999:1-2010:4). This VAR technique provides an indication of how unusual development in the consumer prices inflation was during a given period, and how the contribution of different shocks was over that time period.²⁷

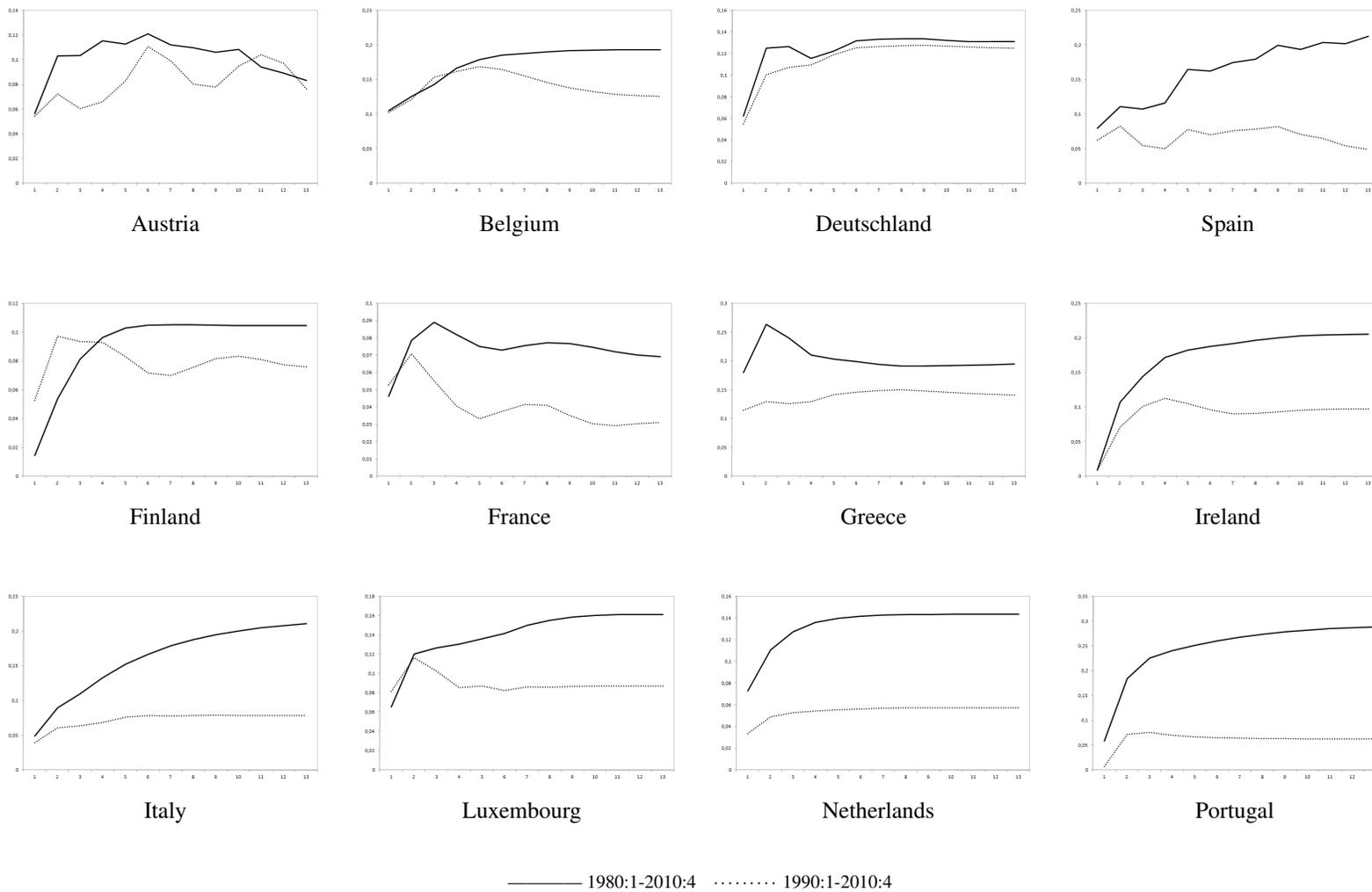
²⁴More details on CANDELON and LUTKEPOHL (2001) chow tests in Appendix C.5.

²⁵For some countries, the structural shift does not happen exactly in 1990:3, but in the vicinity of that date.

²⁶Since the European sovereign debt crisis, the term "GIPS" is used to refer to this group of countries as a label for heavily-indebted economies.

²⁷We talk about "inflation" of consumer prices since our CVAR model is estimated in log differences.

Figure 3.7: Comparison of response of consumer prices



Historical decompositions was employed by HAHN (2003) to assess the contribution of external shocks, occurred since the start of the EMU in January 1999, to inflation at different price stages. According to the author, since the start of the EMU in 1999, oil prices and exchange rate shocks seem to have contributed strongly to increase inflation in the euro area. However, HAHN (2003) focused on the first four years of the monetary union (from 1999 to 2002), given that the time horizon since the introduction of the euro is rather short. Thus, in our study, we compute the historical decompositions for a larger sample period, i.e. during the three stages of EMU.

To compute the contribution of the respective shocks on consumer prices inflation over the time period of interest, we will proceed as follows: first, we consider the actual development of consumer prices inflation series (second column in Table 3.13); second, a base projection is made using the actual data of consumer prices until 1989:4 and assuming no subsequent shocks occur in any of the variables of the model after 1989:4 (column three in Table 3.13); next, we compute the projection error as the difference between the actual development and the projected development (forth column in Table 3.13). Finally, the projection error can be decomposed into the contributions from the respective shocks to consumer prices variation. Given that projection error gives the contribution of all shocks that occurred over a time period on consumer prices, the contribution of one shock is then derived as the difference between the projection including this shock of interest and the projection excluding all shocks (columns five to last in Table 3.13).

In Table 3.13, we report the results of the historical decompositions which represent the average over each sub-period. We note also that we combine shocks into four groups: demand and supply shocks (oil price and real GDP), external shocks (exchange rate and import price), domestic price shocks (producer prices (PPI) and consumer prices (CPI), and interest shocks, i.e. monetary shocks. Now, turning to the analysis of the contribution of different shocks of the consumer prices inflation. Beginning with the sub-period of 1990-1998, we observe that actual consumer prices inflation was above its projection in about half of EA countries, namely in Austria, Spain, France, Greece, Netherlands. For example, in Spain, the shocks occurred since 1989 contributed to increase consumer price inflation by 0.08 percentage points. The external shocks (exchange rate and import prices) were a slightly positive contributor to inflation

in this country, they accounted for about 0.01 percentage points of consumer price inflation. As regards the other shocks, the most inflationary impacts is due to domestic prices shocks (producer and consumer prices together), accounting for 0.06 percentage points. On the other hand, for the rest of EA countries (except Ireland), consumer price inflation was exceptionally low during 1990-1998, as inflation rates was below the base projection, namely in Belgium, Germany, Finland, Italy, Luxembourg and Portugal. For the latter country, actual consumer prices inflation was 0.28 percentage points below its projection on average during 1990-1998. The external shocks contributed strongly to lowering consumer prices inflation by having a disinflationary impact of -0.17%. In addition, monetary shocks was also an important negative contributors to inflation reduction (-0,09 percentage points), suggesting that Portugal may have conducted a tighter monetary policy in anticipation of the creation of the euro.

As regards the sub-sample of 1999-2010, it is interesting to note that for countries where the inflation was low during 1990-1998, now, the actual consumer prices inflation is above the model's base projection. This is true for more than half of our EA sample. This may due to the inflationary effects of large depreciation of the euro in the first three years of his existence. In Belgium, for example, the actual consumer price inflation was (negatively) close to its projection before 1999, while during the third stage of EMU, the actual inflation is higher than projected by 0.05 percentage points, with an inflationary effect of external shocks by 0.02%. Also, it is worth stressing that, contrary to the sub-period 1990-1998, the external shocks was important contributors to consumer prices inflation in most of EA members during 1999-2010. Especially, in four EA countries, namely Germany, Finland and Netherlands, external factors have the relative higher inflationary effect compared to the other shocks (accounted for about 0.03% of consumer prices inflation). As mentioned before, the euro zone has experienced a large depreciation of the single currency since its creation in 1999. Between the end of 1998 to the last quarter of 2001, the euro depreciated by nearly 20% in nominal effective terms. Thereby, this decline in the value of the euro may explain the rise of inflationary pressure since 1999. To sum up, compared to the period of 1990-1998, external factors had an important inflationary impacts on inflation since the starting of the third stage of EMU. This finding is line with HAHN (2003) who found that exchange rate shocks is an important contributor to inflation increase during the first three years of the euro zone.

Table 3.13: Historical decomposition of consumer prices

Country	Actual	Projection	Projection Error	Contribution of Shocks			
				Oil price & GDP	External factors	PPI & CPI	Interest Rate
Austria							
1990 - 1998	0,63	0,55	0,08	0,00	-0,01	0,07	0,01
1999 - 2010	0,47	0,50	-0,03	-0,01	0,02	-0,03	0,00
Belgium							
1990 - 1998	0,52	0,53	-0,01	-0,03	0,00	0,03	-0,02
1999 - 2010	0,56	0,51	0,05	0,03	0,02	0,01	0,00
Germany							
1990 - 1998	0,73	0,77	-0,04	-0,01	-0,01	-0,01	-0,01
1999 - 2010	0,40	0,37	0,03	-0,01	0,03	0,01	-0,01
Spain							
1990 - 1998	1,11	1,03	0,08	0,01	0,01	0,06	0,01
1999 - 2010	0,74	0,73	0,02	-0,01	0,01	0,01	0,00
Finland							
1990 - 1998	0,53	0,54	-0,01	-0,01	-0,02	0,02	0,01
1999 - 2010	0,45	0,42	0,03	0,00	0,03	0,00	0,00
France							
1990 - 1998	0,44	0,43	0,01	-0,01	-0,02	0,04	-0,01
1999 - 2010	0,43	0,44	-0,01	-0,01	0,00	0,00	0,00
Greece							
1990 - 1998	2,12	2,07	0,05	0,00	0,01	0,01	0,03
1999 - 2010	0,78	0,84	-0,06	0,01	-0,02	-0,04	-0,01
Ireland							
1990 - 1998	0,55	0,55	0,00	0,00	-0,01	0,00	0,01
1999 - 2010	0,05	0,00	0,08	0,01	0,00	0,07	0,00
Italy							
1990 - 1998	0,96	1,02	-0,06	-0,01	0,00	-0,03	-0,02
1999 - 2010	0,57	0,56	0,01	0,00	0,01	0,00	-0,01
Luxembourg							
1990 - 1998	0,53	0,67	-0,14	0,01	0,01	-0,03	-0,13
1999 - 2010	0,57	0,58	0,00	-0,02	0,01	-0,01	0,01
Netherlands							
1990 - 1998	0,61	0,57	0,04	0,01	-0,07	0,10	0,00
1999 - 2010	0,54	0,54	0,01	0,00	0,03	-0,01	-0,01
Portugal							
1990 - 1998	1,54	1,82	-0,28	-0,02	-0,17	0,00	-0,09
1999 - 2010	0,65	0,67	-0,02	0,00	0,04	-0,04	-0,02

Note: Numbers are the average over each sub-period (expressed in percentage). Actual corresponds to the actual development of the consumer prices. Projected is made using the actual data of consumer prices up to 1989:4 and assuming no subsequent shocks occur in any of the variables of the model after 1998:4. The projection error is defined as the difference between the actual development and the projected development. The contribution of the shock is defined as the difference between the projection including the respective of interest and the projection excluding all shocks.

7. Conclusion

In this chapter, the pass-through of exchange rate to consumer prices was analyzed for 12 EA countries within a CVAR framework. Using quarterly data ranging from 1980:1 to 2010:4, our study provides new up-to-date estimates of ERPT with paying attention to either the time-series properties of data and variables endogeneity. Using the Johansen cointegration procedure, results indicate the existence of one cointegrating vectors at least for each EA country of our sample. When measuring the long-run effect of exchange rate changes on consumer prices, we found a wide dispersion of ERPT rates across countries. The degree of ERPT appears to be most prevalent in Portugal and Greece. For Portugal, a 1% depreciation of exchange rate increases domestic consumer prices by roughly 0.84%, while for Greece, consumer prices rise by 58% following one percent depreciation of exchange rate. While the lowest coefficients of long-run ERPT were found in Germany, Finland and France (not exceeding 0.20%). It is important to note that the higher pass-through coefficients in Greece and Portugal were confirmed in the empirical literature (see GAGNON and IHRIG (2004)). Besides, when assessing the adjustment coefficients, we point out a very slow adjustment of consumer prices towards their long-run equilibrium is found across EA countries. This would explain the weakness of ERPT estimates in the short-run.

As a natural progression from the cointegration analysis, we carried out impulse response functions analysis. This exercise is done using an extended CVAR model that incorporates features of a distribution chain pricing framework in the spirit of MCCARTHY (2007) and HÜFNER and SCHRÖDER (2002). The *CVAR pricing chain model* enables us to examine the pass-through at different stages along the distribution chain, i.e. import prices, producer prices and consumer prices. Our results show a higher pass-through to import prices with a complete pass-through detected in roughly half EA countries after one year. These results are relatively large compared to single-equation literature. The magnitude of the pass-through of exchange rate shocks decline along the distribution chain of pricing with the modest effect is recorded with consumer prices. Also, referring to the magnitude of the pass-through, we can say that results from impulse response functions corroborate in some extent what we find in the cointegration analysis. The highest ERPT estimates were found in Greece and Portugal, while the

lowest was recorded in France and Finland. When assessing possible reasons for cross-country differences in the ERPT, inflation level, inflation volatility and exchange rate persistence are the main macroeconomic factors that influencing the degree of pass-through almost along the distribution pricing chain. The exchange rate volatility is surprisingly negatively correlated with response of different prices index (as in WEBBER (1999) and BARHOUMI (2006)).

Next, we have investigated the contribution of external shocks (exchange rate and import prices shocks together) using the variance decompositions. Results show that external factors explain only a modest proportion of the forecast variance of domestic consumer prices over 1980-2010, while this contribution is to some extent high in Portugal and Greece. For the latter countries, this would explain why ERPT to consumer is relatively large compared to the other EA members. Mainly, three macroeconomic factors - inflation level, inflation volatility and exchange rate persistence - are found to be crucial in explaining the cross-country differences regarding the influence of external shocks. The degree of openness is surprisingly negatively correlated with the contribution of external shocks (see ROMER, 1993). Thereafter, we have tested for the decline of the response of consumer prices across EA countries. According to multivariate time series Chow test, the stability of ERPT coefficients was rejected, and the impulse responses of consumer prices over 1990-2010 provide an evidence of general decline in rates of pass-through in most of EA countries. Finally, using the historical decompositions, we point out that external factors had important inflationary impacts on inflation since 1999, compared to the pre-euro period. This finding is line with HAHN (2003) who found that exchange rate shocks is an important contributor to inflation increase during the first three years of the monetary union.

Appendix C

C.1. Unit root tests

Table C.1: Results of the Unit Root Tests

Country	ADF		KPSS		DF-GLS	
	Level	1 st diff.	Level	1 st diff.	Level	1 st diff.
CPI						
Austria	-2,512	-3,0937*	0,518243**	1,039	-1,154	-3,011*
Belgium	-0,972	-4,1361**	0,414952**	0,120	-1,246	-2,820*
Germany	-1,372	-3,1581*	0,267249**	0,085	-1,843	-3,616**
Spain	-1,667	-4,5183**	0,255063**	0,104	-0,996	-2,757*
Finland	-0,923	-3,4638**	0,556912**	0,069	-1,165	-3,239*
France	-2,084	-5,0115**	0,474786**	0,298	-1,619	-2,038*
Greece	-1,515	-3,051*	0,631763**	0,124	-1,445	-2,894*
Ireland	-1,084	-1,115	0,306229**	0,166	-1,070	-2,126*
Italy	-2,125	-3,1928*	0,547267**	0,086	-1,565	-2,864**
Luxembourg	-1,062	-4,9549**	0,329039**	0,160	-1,693	-1,404
Netherlands	0,192	-4,5356**	0,251542**	0,491894*	-1,436	-2,966*
Portugal	-1,796	-4,6488**	0,304176**	0,558627*	-1,505	-2,245
Nominal Effective Exchange Rate						
Austria	-0,951	-8,5404**	0,561378**	0,373	-1,244	-4,692**
Belgium	-2,931	-7,0648**	0,187107*	0,175	-1,700	-2,953**
Germany	-2,129	-8,4702**	0,393234**	0,086	-2,535	-4,784**
Spain	-3,296	-7,4751**	0,200878*	0,384	-1,121	-3,282**
Finland	-2,243	-7,7206**	0,326954**	0,066	-2,829	-4,158**
France	-1,953	-8,9225**	0,210352*	0,337	-1,466	-3,296**
Greece	-0,771	-8,3779**	0,619331**	0,207	-0,529	-2,045*
Ireland	-2,027	-8,3887**	0,214021*	0,203	-1,400	-2,801**
Italy	-1,904	-7,5047**	0,349329**	0,102	-1,266	-4,457**
Luxembourg	-1,763	-7,2815**	0,289906**	0,113	-1,683	-3,355**
Netherlands	-1,916	-8,0551**	0,227835**	0,120	-2,459	-4,964**
Portugal	-1,942	-5,8794**	0,352049**	0,143	-1,383	-2,469
GDP						
Austria	-2,913	-7,9166**	1,810626**	0,121	-2,158	-2,816**
Belgium	-0,334	-5,0765**	2,499921**	0,130	-1,930	-3,671**
Germany	-1,198	-7,6517**	2,139009**	0,082	-2,602	-4,124**
Spain	-0,979	-2,9660*	2,401370**	0,154	-2,238	-2,844**
Finland	-0,418	-8,5876**	2,276504**	0,098	-2,557	-2,466*
France	-1,215	-4,8021**	2,410361**	0,163	-2,090	-2,495*
Greece	0,108	-3,9337**	2,262626**	0,368	-0,792	-4,126**
Ireland	-0,717	-3,1905*	2,405650**	0,419	-1,583	-2,658**
Italy	-1,873	-6,9938**	2,355765**	0,305	-0,941	-3,794**
Luxembourg	-0,723	-10,6741**	2,409973**	0,166	-1,799	-3,460**
Netherlands	-0,199	-10,0595**	2,427777**	0,288	-1,507	-2,720**
Portugal	-1,194	-3,7643**	2,376592**	0,395	-1,911	-2,143*

Note: The tests were performed on the logs of the series (except interest rates) for levels including an intercept and trend. The critical values at 1% and 5% levels respectively are: ADF: -3.99, -3.43; KPSS: 0.216, 0.146; DFGLS: -3.48, -2.89. For the first-differences, the tests included only an intercept and were based on the following critical values at the 1%, 5%, and 10% levels respectively: ADF: -3.46, -2.88; KPSS: 0.739, 0.463; DFGLS: -2.58, -1.95. ** and * respectively refer to significance at the 1% and 5%.

Table C.1: *Continued*

Country	ADF		KPSS		DF-GLS	
	Level	1 st diff.	Level	1 st diff.	Level	1 st diff.
Interest Rate						
Austria	-3,129	-6,3400**	1,574195**	0,084	-1,568	-2,144*
Belgium	-2,081	-6,2442 **	0,223654**	0,054	0,271	-3,309**
Germany	-1,957	-5,3625 **	1,476766**	0,096	-1,513	-4,384**
Spain	-1,405	-5,7984 **	2,354212**	0,092	-0,032	-4,463**
Finland	-0,988	-7,4761**	2,146838**	0,091	-0,770	-4,944**
France	-1,054	-6,3973 **	2,201712**	0,067	-0,100	-4,562**
Greece	0,321	-4,7748**	2,246433**	0,227	-0,166	-4,951**
Ireland	-1,857	-7,7125**	2,214257**	0,051	0,002	-4,256**
Italy	-0,957	-5,6865**	2,371547**	0,115	0,082	-2,107*
Luxembourg	-2,081	-6,2442**	2,156896**	0,054	0,271	-3,309**
Netherlands	-2,028	-6,1353**	1,681538**	0,054	-0,222	-5,270**
Portugal	-1,054	-5,6402**	2,262505**	0,109	-0,616	-4,737**
Import Prices Index						
Austria	-1,740	-5,2184**	0,293895**	0,089	-2,050	-4,303**
Belgium	-2,976	-4,9145**	0,171710*	0,293	-1,370	-3,852**
Germany	-2,448	-6,7764**	0,157448*	0,155	-2,178	-2,000*
Spain	-0,596	-5,9655**	0,181320*	0,459	-1,552	-3,384**
Finland	-2,475	-8,9163**	0,316630**	0,416	-1,271	-4,284**
France	-1,908	-5,7997**	0,169495*	0,193	-1,480	-3,553**
Greece	-0,600	-9,2081**	0,188599*	0,152	-1,870	-4,641**
Ireland	-1,404	-9,2233**	0,162430*	0,078	-2,019	-2,045*
Italy	-0,821	-6,2556**	2,070254**	0,393	-1,560	-2,495*
Luxembourg	-0,360	-9,0961**	0,095	0,039	-1,811	-3,822**
Netherlands	-0,180	-6,6482**	0,318671**	0,167	-2,118	-4,411**
Portugal	-0,983	-3,3908*	0,478687**	0,052	-0,501	-3,020*
PPI						
Austria	-1,482	-4,6157**	0,385971**	0,126	-2,207	-4,501**
Belgium	-2,677	-5,8743**	0,294241**	0,259	-1,799	-4,331**
Germany	-3,003	-6,6539 **	0,156822*	0,054	-1,578	-4,004**
Spain	-2,049	-3,5715**	0,322361**	0,049	-1,150	-3,064*
Finland	-2,657	-5,5032 **	0,265933**	0,067	-1,177	-4,338**
France	-2,886	-3,3363**	0,561824**	0,420	-1,932	-3,530**
Greece	-3,086	-6,3399**	0,295182**	0,059	-2,648	-3,482**
Ireland	-3,056	-5,5857**	0,344062**	0,227	-0,610	-0,812
Italy	-2,400	-3,3184*	0,533521**	0,101	-2,281	-3,842**
Luxembourg	-3,060	-6,0600**	0,350788**	0,125	-1,232	-4,713**
Netherlands	-2,017	-5,4149**	0,289999**	0,114	-1,690	-4,385**
Portugal	-3,126	-6,3424**	0,338455**	0,400	-1,083	-1,338
Foreign costs						
Austria	-1,902	-8,9765**	0,431481**	0,090	-1,871	-5,005**
Belgium	-3,247	-7,3116**	0,118	0,087	-1,582	-3,085**
Germany	-1,917	-9,2879 **	0,348081**	0,072	-2,476	-4,254**
Spain	-2,518	-7,1836**	0,140	0,253	-1,681	-3,368**
Finland	-1,950	-8,0770**	0,321281**	0,302	-2,398	-3,713**
France	-3,376	-8,0228 **	0,128	0,095	-1,466	-3,866**
Greece	-0,771	-9,1764**	0,632549**	1,177470**	-0,382	-2,045*
Ireland	-2,377	-9,2779**	0,383129**	0,306	-1,089	-4,729**
Italy	-1,873	-7,9844**	0,465492**	0,430	-0,922	-4,490**
Luxembourg	-3,060	-8,0464**	0,113	0,076	-2,093	-3,279**
Netherlands	-1,645	-8,9249**	1,443464**	0,090	-2,606	-4,025**
Portugal	-2,045	-6,4277**	0,540070**	0,251	-1,332	-2,501
Oil Price Index						
	-1,778	-9,4680**	0,545785**	0,200	-1,111	-3,199**

Note: The tests were performed on the logs of the series (except interest rates) for levels including an intercept and trend. The critical values at 1% and 5% levels respectively are: ADF: -3.99, -3.43; KPSS: 0.216, 0.146; DFGLS: -3.48, -2.89. For the first-differences, the tests included only an intercept and were based on the following critical values at the 1%, 5%, and 10% levels respectively: ADF: -3.46, -2.88; KPSS: 0.739, 0.463; DFGLS: -2.58, -1.95. ** and * respectively refer to significance at the 1% and 5%.

C.2. Akaike Information Criterion (AIC) for Lag selection

Table C.2: Lag selection for baseline VECM

<i>Lag Order</i>	Austria	Belgium	Germany	Spain	Finland	France
0	-707,649	-1719,496	-1802,515	-981,602	-91,776	-1116,134
1	-3116,696	-4064,421	-4674,068	-3685,959	-2301,892	-3608,798
2	-3167,887	-4098,574	-4704,113	-3721,812	-2297,081	-3643,728
3	-3140,334	-4050,572	-4697,863	-3741,014	-2294,245	-3627,965
4	-3119,215	-3937,324	-4654,755	-3684,688	-2294,572	-3603,182
5	-3159,858	-3818,671	-4575,063	-3637,260	-2288,652	-3587,996
6	-3142,221	-3660,384	-4455,447	-3543,071	-2237,942	-3532,887
7	-3095,840	-3463,321	-4295,768	-3401,457	-2179,878	-3483,165
8	-3047,109	-3224,676	-4112,951	-3182,100	-2126,362	-3414,625
<i>Lag Order</i>	Greece	Ireland	Italy	Luxembourg	Netherlands	Portugal
0	45,165	-74,602	-328,235	-338,479	-555,000	52,303
1	-1992,475	-2052,221	-2647,038	-2441,852	-2654,382	-2381,231
2	-2035,218	-2076,474	-2683,614	-2445,347	-2681,073	-2424,304
3	-2081,116	-2085,094	-2676,574	-2457,898	-2657,344	-2417,389
4	-2055,176	-2053,489	-2653,193	-2419,032	-2631,082	-2387,883
5	-2088,453	-1996,437	-2614,111	-2372,169	-2624,315	-2327,063
6	-2052,830	-1937,264	-2570,460	-2321,386	-2584,743	-2275,389
7	-2002,637	-1866,812	-2521,844	-2264,651	-2537,194	-2213,436
8	-1926,744	-1840,338	-2471,435	-2205,522	-2469,914	-2171,147

Note: The minimum of the AIC values are in bold.

Table C.3: Lag selection for pricing chain VECM

<i>Lag Order</i>	Austria	Belgium	Germany	Spain	Finland	France
0	-4086,440	-4265,487	-4849,386	-4073,256	-4136,728	-4261,113
1	-4197,367	-4422,103	-4985,606	-4194,432	-4242,746	-4413,646
2	-4337,963	-4425,441	-5016,705	-4292,703	-4221,732	-4425,795
3	-4329,941	-4320,672	-4967,432	-4241,651	-4209,630	-4361,654
4	-4344,365	-4236,356	-4954,561	-4216,806	-4186,176	-4340,113
5	-4251,831	-4094,537	-4855,865	-4144,285	-4071,849	-4210,641
6	-4158,885	-3946,675	-4717,711	-4037,858	-3954,318	-4061,363
7	-4011,018	-3748,484	-4551,084	-3858,164	-3786,388	-3886,029
8	-3800,203	-3451,027	-4368,828	-3694,118	-3551,730	-3660,326
<i>Lag Order</i>	Greece	Ireland	Italy	Luxembourg	Netherlands	Portugal
0	-3331,906	-4218,482	-3941,860	-3620,628	-4345,221	-4247,504
1	-3468,345	-4326,111	-4133,021	-3686,347	-4448,917	-4790,432
2	-3513,692	-4340,178	-4120,329	-3711,246	-4431,615	-4773,790
3	-3484,442	-4316,416	-4120,570	-3644,666	-4380,499	-4729,041
4	-3440,340	-4257,068	-4059,120	-3582,060	-4382,978	-4687,828
5	-3358,258	-4214,308	-4024,986	-3510,567	-4324,392	-4626,802
6	-3223,859	-4125,983	-3917,337	-3435,711	-4252,896	-4566,572
7	-2968,743	-4047,076	-3799,292	-3325,171	-4149,634	-4485,263
8	-2694,974	-3941,302	-3662,840	-3160,203	-4039,168	-4372,393

Note: The minimum of the AIC values are in bold.

C.3. LR Trace Test Results

Table C.4: Johansen Trace Test

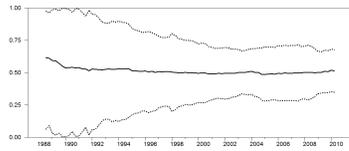
$H_0: \text{rank} = r$	Austria	Belgium	Germany	Spain	Finland	France
0	98,540 (0,000)	126,785 (0,000)	150,707 (0,000)	211,569 (0,000)	217,555 (0,000)	165,362 (0,000)
1	41,959 (0,160)	59,014 (0,003)	76,514 (0,002)	108,246 (0,000)	50,993 (0,023)	96,584 (0,000)
2	13,324 (0,875)	23,376 (0,236)	43,402 (0,043)	24,111 (0,460)	19,720 (0,453)	51,757 (0,006)
3	3,048 (0,957)	6,922 (0,593)	13,679 (0,688)	12,779 (0,390)	6,436 (0,649)	21,500 (0,173)
4	0,411 (0,521)	2,339 (0,126)	2,708 (0,896)	3,425 (0,515)	0,061 (0,805)	7,596 (0,306)
$H_0: \text{rank} = r$	Greece	Ireland	Italy	Luxembourg	Netherlands	Portugal
0	135,823 (0,000)	114,898 (0,000)	125,457 (0,000)	81,370 (0,004)	113,458 (0,001)	164,218 (0,000)
1	85,627 (0,007)	62,078 (0,001)	74,453 (0,004)	47,904 (0,048)	57,245 (0,314)	93,617 (0,000)
2	47,398 (0,125)	28,905 (0,064)	43,141 (0,046)	21,175 (0,357)	26,916 (0,794)	50,983 (0,005)
3	25,689 (0,236)	5,763 (0,725)	22,141 (0,137)	4,666 (0,840)	12,708 (0,784)	19,253 (0,272)
4	12,220 (0,172)	0,073 (0,787)	4,760 (0,636)	1,932 (0,165)	0,060 (1,000)	2,990 (0,867)

Note: p -value are in parentheses.

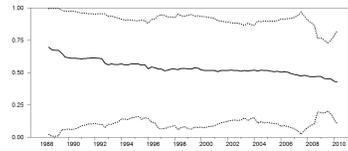
C.4. Recursive Analysis of Eigenvalues

A variety of diagnostic tools can be used to investigate parameter constancy by means of recursive estimation as proposed by HANSEN and JOHANSEN (1999). Starting from a base sample X_{-k+1}, \dots, X_{T_0} , the eigenvalues are calculated recursively for increasing samples X_{-k+1}, \dots, X_t for $t = T_0 + 1, \dots, T$ based upon which the diagnostic tests are calculated. In Figure C.1, We report the plots of time paths the largest eigenvalue of the unrestricted VAR model for each country. Non-constancy of β_i or α_i will be reflected in the eigenvalue λ_i .

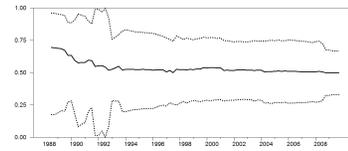
Figure C.1: Time paths of eigen values with 95% confidence bands



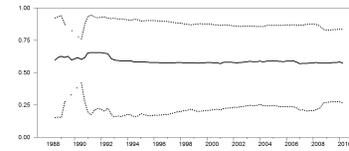
Austria



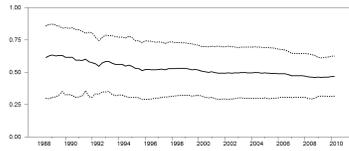
Belgium



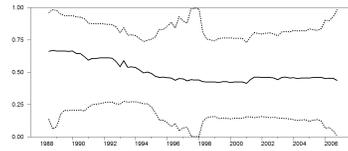
Deutschland



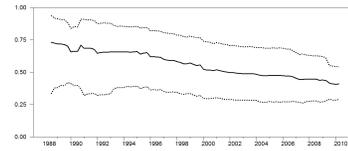
Spain



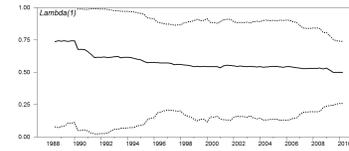
Finland



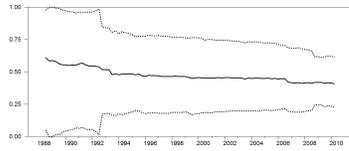
France



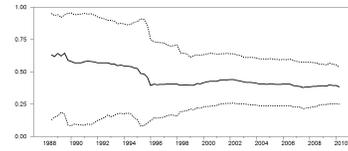
Greece



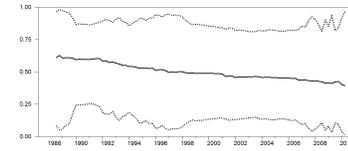
Ireland



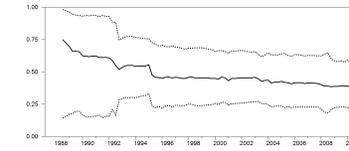
Italy



Luxembourg



Netherlands



Portugal

C.5. Chow tests for multiple time series systems

CANDELON and LUTKEPOHL (2001) consider two versions of Chow tests, sample-split (SS) tests and break-point (BP) tests. The BP Chow test for checking for a structural break in period T_B proceeds as follows. The model under consideration is estimated from the full sample of T observations and from the first T_1 and the last T_2 observations, where $T_1 < T_B$ and $T_2 = T - T_B$. The resulting residuals are denoted by \hat{u}_t , $\hat{u}_t^{(1)}$ and $\hat{u}_t^{(2)}$, respectively. Using the notation $\tilde{\Sigma}_u = T^{-1} \sum_{t=1}^T \hat{u}_t \hat{u}_t'$, $\tilde{\Sigma}_{1,2} = (T_1 + T_2)^{-1} (\sum_{t=1}^{T_1} \hat{u}_t \hat{u}_t' + \sum_{t=T-T_2+1}^T \hat{u}_t \hat{u}_t')$, $\tilde{\Sigma}_{(1)} = T_1^{-1} \sum_{t=1}^{T_1} \hat{u}_t^{(1)} \hat{u}_t^{(1)'}$ and $\tilde{\Sigma}_{(2)} = T_2^{-1} \sum_{t=T-T_2+1}^T \hat{u}_t^{(2)} \hat{u}_t^{(2)'}$, the BP test statistic is:

$$\lambda_{BP} = (T_1 + T_2) \log \det \tilde{\Sigma}_{1,2} - T_1 \log \det \tilde{\Sigma}_{(1)} - T_2 \log \det \tilde{\Sigma}_{(2)} \approx \chi^2(k), \quad (\text{C.1})$$

Where k is the difference between the sum of the number of parameters estimated in the first and last subperiods and the number of parameters in the full sample model. Note that also the potentially different parameters in the white noise covariance matrix are counted. The null hypothesis of constant parameters is rejected if λ_{BP} is large.

$$\lambda_{SS} = (T_1 + T_2) [\log \det \tilde{\Sigma}_{1,2} - \log \det (T_1 + T_2)^{-1} (T_1 \tilde{\Sigma}_{(1)} + T_2 \tilde{\Sigma}_{(2)})] \approx \chi^2(k^-) \quad (\text{C.2})$$

Here k^- is the difference between the sum of the number of coefficients estimated in the first and last subperiods and the number of coefficients in the full sample model, not counting the parameters in the white noise covariance matrix.

CANDELON and LUTKEPOHL (2001) pointed out that especially for multivariate time series models the asymptotic χ^2 -distribution may be a poor guide for small sample inference. Even adjustments Based on F approximations can lead to distorted test sizes. Therefore, they have proposed using bootstrap versions of the Chow tests to improve their small sample properties. They are computed as follows. From the estimation residuals \hat{u}_t , centered residuals $\hat{u}_1 - \bar{\hat{u}}, \dots, \hat{u}_T - \bar{\hat{u}}$ are computed. Bootstrap residuals

u_1^*, \dots, u_T^* are generated by randomly drawing with replacement from the centered residuals. Based on these quantities, bootstrap time series are calculated recursively starting from given pre-sample values y_{p+1}, \dots, y_0 . Then the model is reestimated with and without allowing for a break and bootstrap versions of the statistics of interest, say λ_{BP}^* and λ_{SS}^* are computed. The p -values of the tests are estimated as the proportions of values of the bootstrap statistics exceeding the corresponding test statistic based on the original sample.

Table C.5: Chow test for VECM

Chow Test	Austria	Belgium	Germany	Spain	Finland	France
Break Point test	741,118	1038,114	590,512	783,370	550,273	928,758
<i>bootstrapped p-value</i>	(0,000)	(0,020)	(0,000)	(0,000)	(0,000)	(0,070)
Sample Split test	481,848	460,226	419,039	495,692	402,790	639,689
<i>bootstrapped p-value</i>	(0,000)	(0,010)	(0,020)	(0,000)	(0,010)	(0,000)
Chow Test	Greece	Ireland	Italy	Luxembourg	Netherlands	Portugal
Break Point test	896,897	663,445	414,849	333.706	755,095	759,694
<i>bootstrapped p-value</i>	(0,100)	(0,000)	(0,000)	(0,030)	(0,000)	(0,000)
Sample Split test	576,060	367,118	410.522	214.397	467,483	448,913
<i>bootstrapped p-value</i>	(0,000)	(0,090)	(0,000)	(0,040)	(0,000)	(0,000)

Note: Bootstrapped p -values are obtained from 1000 bootstrap replication.

C.6. Generalized impulse response for consumer prices

Based on KOOP, PESARAN, and POTTER (1996), PESARAN and SHIN (1998) proposed the generalized impulse response where ordering does not matter. The generalized impulse response function of x_t at horizon n is defined by:

$$GI_x(n, \delta, \Omega_{t-1}) = E(x_{t+n} | \varepsilon_t = \delta, \Omega_{t-1}) - E(x_{t+n} | \Omega_{t-1}) \quad (\text{C.3})$$

Where Ω_{t-1} a non-decreasing information set, denotes the known history of the economy up to time $t - 1$ and $\delta = (\delta_1, \dots, \delta_m)'$ is some hypothetical m vector of shocks hitting the economy at time t . Considering the following the infinite moving average

representation of a VAR system:

$$x_t = \sum_{i=0}^{\infty} A_i \varepsilon_{t-i} \quad t = 1, \dots, T, \quad (\text{C.4})$$

Using (C.3) in (C.4) gives:

$$GI_x(n, \delta, \Omega_{t-1}) = A_n \delta \quad (\text{C.5})$$

Which is independent of Ω_{t-1} but depends on the composition of shocks defined by δ . The appropriate choice of hypothesized vector of shocks, δ , is central to the properties of the impulse response function. The traditional approach, suggested by SIMS (1980), is to resolve the problem surrounding the choice of δ by using the Cholesky decomposition of $\Sigma = E(\varepsilon_t \varepsilon_t')$, the variance-covariance matrix of ε_t :

$$PP' = \Sigma \quad (\text{C.6})$$

Where P is an $m \times m$ lower triangular matrix. Then, (C.4) can be rewritten as:

$$x_t = \sum_{i=0}^{\infty} (A_i P)(P^{-1} \varepsilon_{t-i}) = \sum_{i=0}^{\infty} (A_i P) \xi_{t-i} \quad t = 1, \dots, T, \quad (\text{C.7})$$

such that $\xi_t = P^{-1} \varepsilon_t$ are orthogonalized; namely, $E(\xi_t \xi_t') = I_n$. Hence, the $n \times 1$ vector of the orthogonalized impulse response function of a unit shock to the j th equation on x_{t+1} is given by

$$\psi_j^0(n) = A_n P e_j, \quad n = 1, 2, \dots \quad (\text{C.8})$$

Where e_j is an m selection vector with unity as its j th element and zeros elsewhere.

An alternative approach would be to use (C.3) directly, but instead of shocking all the elements of ε_t , to shock only one element, say its j th element, and integrate out the effects of other shocks using an assumed or the historically observed distribution of the errors. In this case one would have

$$GI_x(n, \delta_j, \Omega_{t-1}) = E(x_{t+n} | \varepsilon_{jt} = \delta_j, \Omega_{t-1}) - E(x_{t+n} | \Omega_{t-1}) \quad (\text{C.9})$$

Assuming that ε_t has a multivariate normal distribution, KOOP, PESARAN, and POTTER (1996) show that:

$$E(\varepsilon_t | \varepsilon_{jt} = \delta_j) = (\sigma_{1j}, \dots, \sigma_{mj})' \sigma_{jj}^{-1} \delta_j = \Sigma e_j \sigma_{jj}^{-1} \delta_j \quad (\text{C.10})$$

Hence, the $m \times 1$ vector of the (unscaled) generalized impulse response of the effect of a shock in the j th equation at time t on x_{t+1} is given by:

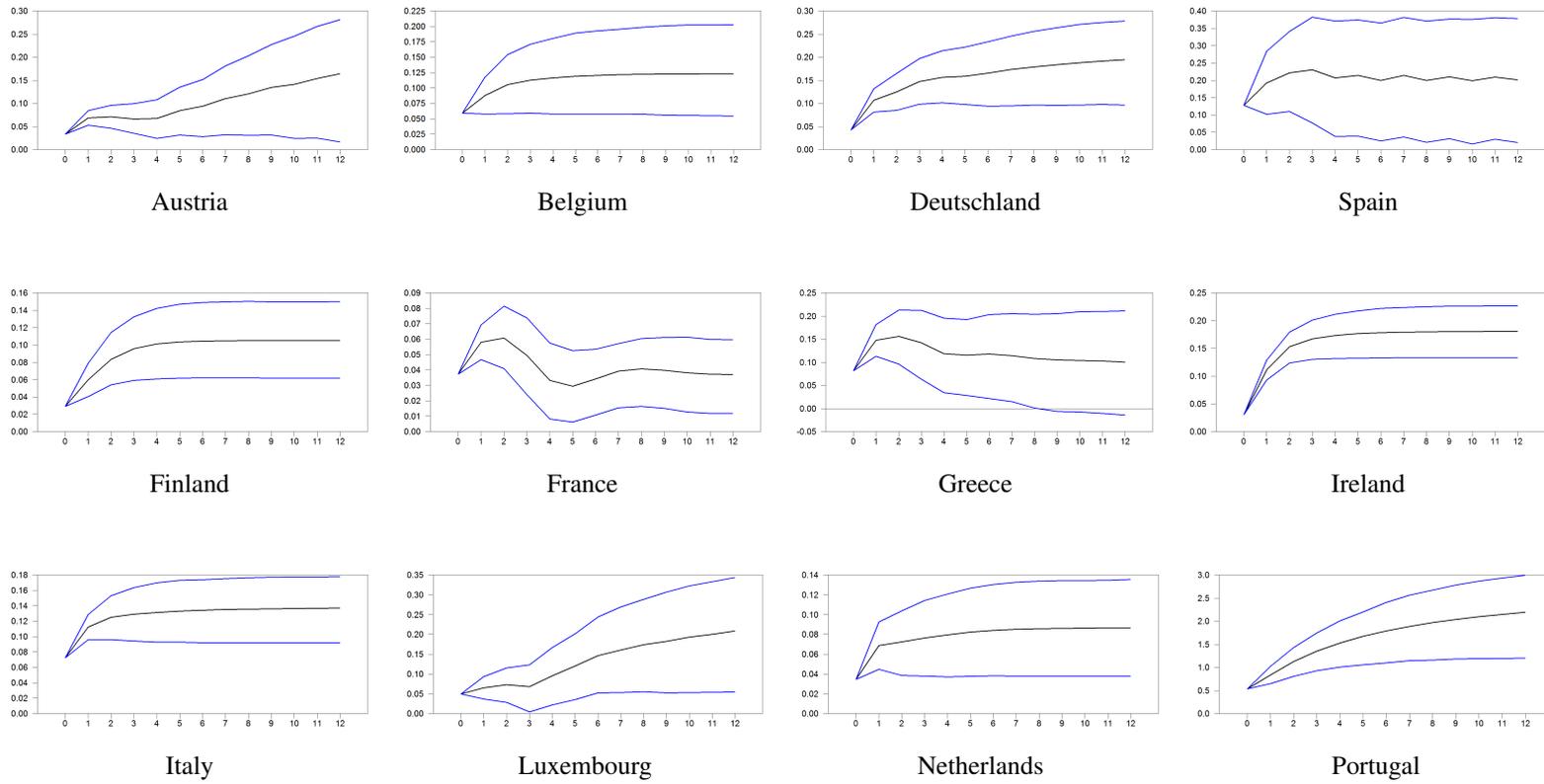
$$\left(\frac{A_n \Sigma e_j}{\sqrt{\sigma_{jj}}} \right) \left(\frac{\delta_j}{\sqrt{\sigma_{jj}}} \right), \quad n = 1, 2, \dots \quad (\text{C.11})$$

By setting $\delta_j = \sqrt{\sigma_{jj}}$, PESARAN and SHIN (1998) derived the scaled generalized impulse response function by:

$$\psi_j^g(n) = \sigma_{jj}^{\frac{1}{2}} A_n \Sigma e_j, \quad n = 1, 2, \dots \quad (\text{C.12})$$

This latter measures the effect of one standard error shock to the j th equation at time t on expected values of x at time $t + n$.

Figure C.2: Response of consumer prices to 1% exchange rate shock



Chapter 4

Nonlinear Mechanisms of Exchange Rate Pass-through

For information:

Section 6.2.1 and Section 6.2.2 are published, respectively, in the following refereed journals:

- Nidhaleddine Ben Cheikh, (2012) "Non-linearities in exchange rate pass-through: Evidence from smooth transition models", *Economics Bulletin*, Vol. 32 (3), 2530-2545.
- Nidhaleddine Ben Cheikh (2012) "Asymmetric Exchange Rate Pass-Through in the Euro Area: New Evidence from Smooth Transition Models", *Economics: The Open-Access, Open-Assessment E-Journal*, Vol. 6(39), 1-28.

1. Introduction

The issue of asymmetries and nonlinearities is one of the burgeoning topics in the literature of Exchange Rate Pass-Through (ERPT). In fact, there are various circumstances that could generate asymmetric adjustment of prices to exchange rate changes which can't be modeled within a simple linear framework. Some spectacular exchange rate movements like those experienced by the US dollar in the 1980s seems to be an illustration of an asymmetric pattern. The appreciation of the dollar against the Deutsch mark amounted to 70% between 1980-1985 and the subsequent depreciation amounted to 80% by the end of 1987.¹ Similarly, since the creation of the euro area (EA), there has been a large depreciation of the European currency against the US dollar from 1999 till the last of 2001. Similarly, since the creation of the euro area (EA), there has been a large depreciation of the European currency against the US dollar from 1999 till the last of 2001. After that date, the euro started appreciating to become a strong and well established currency. It is expected that these dramatic exchange rate developments may affect asymmetrically domestic prices, raising the question of the presence of a nonlinear dynamic in ERPT mechanism (see BUSSIÈRE (2012)).

Nonetheless, the empirical literature has paid little attention to the issue of asymmetries and nonlinearities in ERPT in spite of its strong policy relevance. The number of studies which have investigating for nonlinearities in this context is to date relatively scarce, and most of papers assume linearity rather than testing it. The sparse empirical evidence on this area of research has put forth the role of exchange rate movements in generating nonlinearities. According to this literature, mainly, there are two potential sources of pass-through asymmetry. On one hand, asymmetry can arise from the *direction* of exchange rate changes i.e., in response to currency depreciations and appreciations. On the other hand, the extent of pass-through may also respond asymmetrically to the *magnitude* of exchange rate movements, i.e. depending on whether exchange rate changes are large or small. However, as pointed by MARAZZI et al. (2005), previous studies provide mixed results with no clear support for the existence of important nonlinearities. If the existing literature is not conclusive,

¹In spite of these dramatic changes in the value of the dollar during this period, YANG (2007) provide a weak evidence of asymmetric ERPT between appreciation and depreciation.

there are two important caveats should be noted in this regard. First, ERPT is not depending exclusively on exchange rate changes, there are various factors, including macroeconomic variables, which might influence the pass-through mechanisms. Thus, other sources of nonlinearities or asymmetries may exist. For instance, GOLDFAJN and WERLANG (2000) report an asymmetric reaction of the ERPT over the business cycle. Second, an appropriate econometric tool is required. Several empirical studies on asymmetries in ERPT experiment a standard linear model augmented with interactive dummy variables. These added interactive terms would account for appreciation or depreciation episodes as well as for some specific events such as unusual exchange rate developments (See YANG (2007)). For example, in order to capture possible asymmetries in ERPT, COUGHLIN and POLLARD (2004) use threshold dummy variables to distinguish between large and small exchange rate changes. The authors choose an arbitrary threshold value for all US industries equal to 3%. A large exchange rate change is defined as being 3% and above, while a small change is below 3%. However, for more accuracy, the threshold level must be estimated from the data instead of using an arbitrary value. So, a relevant econometric method is required. An alternative methodology is to estimate a nonlinear regime-switching model where a grid search is used to select the appropriate threshold. Amongst this class of models, two popular nonlinear models can be mentioned. First, the so-called threshold regression model where the transition across regimes is abrupt.² Second, the smooth transition regression (STR) model with the transition between states is rather smooth.³ In this chapter, we propose to use the second type of regime-switching model, namely a class of smooth transition regression models, in order to investigate for the presence of nonlinear mechanisms in the ERPT.

Recently, there has been an increasing interest for models with regime-switching behavior in modeling the ERPT, although the number of studies is still sparse. CORREA and MINELLA (2006) estimate a threshold regression model for the Brazilian economy in order to check for possible nonlinearities in ERPT. In addition to exchange rate changes, the authors test for other sources of nonlinearity, namely exchange rate volatility and business cycle. Their results reveal that pass-through is higher when the economy is growing faster, when the exchange rate depreciates above a certain threshold, and when exchange rate volatility is lower. Regarding the STR models, there are very

²The univariate case is known as the threshold autoregressive (TAR) Model.

³The univariate case is known as smooth transition autoregressive (STAR) Model.

few studies that using these non-linear models in the context of pass-through. SHINTANI, TERADA-HAGIWARA, and TOMOYOSHI (2013) estimated the ERPT to US domestic prices with respect to inflation level. They find that the period of low ERPT would be associated with the low inflation environment. In a more complete study, NOGUEIRA JR. and LEON-LEDESMA (2008) examine the possibility of non-linear pass-through for a set of inflation target countries. They found that asymmetric adjustment of prices to exchange rate changes can be related to several macroeconomic factors, including inflation rate, the size of exchange rate changes, two measures of macroeconomic instability and output growth.⁴ In a similar vein, HERZBERG, KAPETANIOS, and PRICE (2003) analyzed the ERPT into UK import prices using a STR model but did not find any evidence of non-linearity.

THE AIM OF this chapter is to fill the gap in empirical evidence on the nonlinearities in ERPT. More precisely, we follow SHINTANI, TERADA-HAGIWARA, and TOMOYOSHI (2013) and NOGUEIRA JR. and LEON-LEDESMA (2008) by using a STR models to estimate the extent of pass-through. We focus on “consumer-price pass-through”, i.e. the sensitivity of consumer prices to exchange rate changes. Unlike the cited studies, we are interested in the EA case since we expect that the different exchange rate arrangements experienced by the monetary union members would generate a nonlinear mechanism in ERPT. To our knowledge, there is no other study has applied a nonlinear STR estimation approach in this context. We note that the presence of nonlinearities is tested with respect to different macroeconomic determinants, namely the inflation environment, the direction and the size of exchange rate changes, the economic activity and the macroeconomic instability.

To preview our results, we found that the degree of pass-through respond non-linearly to the inflation environment, that is, ERPT is higher when the inflation level surpasses some limit. The time-varying ERPT coefficients point out that exchange rate pass-through has declined over time in the EA countries which are due to the shift to a low-inflation environment. When considering the direction of exchange rate change as a potential source of nonlinearities - that is, whether ERPT asymmetrically to appreciation - we report mixed results with no clear-cut evidence about the direction of asymmetry. This is not surprising since, in theory, an appreciation can lead to either a higher or lower

⁴More details on these studies in section 4

rate of pass-through than depreciation. Next, we check the asymmetry of pass-through with respect exchange rate magnitude. We find that large exchange rate changes elicit greater pass-through than small ones. Results give a broad evidence of the presence of menu costs, when exchange rate changes exceed some threshold, firms are willing to pass currency movements through their prices. These findings seem to explain why ERPT was greater during the EMS Crisis and at the launch of the euro. Thereafter, the source of nonlinearities considered in our study is relative to business cycle. Our results provide a strong evidence of nonlinearity in 6 out of 12 EA countries with significant differences in the degree of ERPT between the periods of expansion and recession. However, we find no clear direction in this regime-dependence of pass-through to business cycle. In some countries, ERPT is higher during expansions than in recessions; however, in other countries, this result is reversed. These cross-country differences in the nonlinear mechanism of pass-through would have important implications for the design of monetary policy and the control of inflation in the EA context. Finally, we test whether periods of macroeconomic instability/confidence crisis may alter the extent of pass-through in a nonlinear way. In the light of the recent European sovereign debt crisis, we propose to use 10-year government bond yield differentials (versus Germany) as an indicator of macroeconomic instability. Our estimation is conducted only for the heavily-indebted EA economies i.e. the GIIPS group (Greece, Ireland, Italy, Portugal, and Spain). Our results show that in periods of widening spreads, which corresponds to episodes of confidence crisis, the degree of ERPT is higher.

The remainder of the **Chapter 4** is structured as follows: Section 2 gives the reasons for the potential existence of a non-linear ERPT. In section 3, the modeling strategy of estimating a STR model is presented. Section 4 discusses the existing empirical literature that implemented STR models to measure the pass-through. Section 5 describes data set and the final specification to estimate. Section 6 presents the main empirical results and Section 7 concludes.

2. ERPT and nonlinearities

2.1. Why ERPT could be non-linear?

The empirical literature has paid little attention to the issue of nonlinearities and asymmetry in ERPT in spite of its strong policy relevance. The number of studies which have investigating for nonlinearities in this context is to date relatively scarce, and most of papers assume linearity rather than testing it. From a theoretical point of view, the assumption that ERPT is linear and symmetric is not realistic. In fact, there are various circumstances that could generate asymmetry in the pass-through mechanisms. The sparse empirical evidence on this area of research has put forth the role of exchange rate in generating nonlinearities. According to this literature, potential asymmetric behavior can arise from the *direction* of exchange rate changes i.e., in response to currency depreciations and appreciations (see e.g. MARSTON (1990), GIL-PAREJA (2000) and OLIVEI (2002)). On the other hand, the extent of pass-through may also respond asymmetrically to the size of exchange rate movements, since there is differential effect of large versus small exchange rate changes (COUGHLIN and POLLARD (2004) and BUSSIÈRE (2007)). There is some theoretical (microeconomic) arguments behind the potential asymmetric relationship between the exchange rate and prices. Mainly, we mention three explanations of a possible ERPT asymmetry:

- *Market share objective*: faced with a depreciation of the domestic currency, foreign firms can follow pricing-to-market (PTM) strategy by adjusting their markups to maintain market. However, with an appreciation, they maintain their markups and allow the import price to fall in the currency of destination market. Consequently, the extent of ERPT would be different with respect to exchange rate changes direction. If foreign producers attempt to keep competitiveness and maintain market share, then an appreciation of the domestic currency might cause higher pass-through than a depreciation.
- *Capacity constraints*: quantities may be rigid upwards in the short run. Faced with an appreciation of the importing country's currency, foreign exporters would gain in price competitiveness by passing this exchange rate change into their prices.

But, if foreign firms have already reached full capacity, the ability of increasing sales in destination market is limited, and they may be tempted to increase their markup instead of lowering prices in the currency of the importing country.⁵ As argued by KNETTER (1994), if exporting firms are subject to binding *quantity constraints*, then an appreciation of the currency of the importing country might cause lower pass-through than a depreciation.

It is important to note that the two first arguments have a clear implication for possible nonlinearities in ERPT, but in the same time they give rise to opposite interpretations of asymmetry. According to market share explanation, pass-through will be higher when the importer's currency is appreciating than when it is depreciating. While, the quantity constraint hypothesis suggest the opposite result, and ERPT would be highest when exchange rate is depreciating. Empirically, previous studies provide also no clear-cut evidence on the direction of asymmetry. In some cases the pass-through associated with depreciations exceeded appreciations; however, in other cases, this result is quite the opposite. GIL-PAREJA (2000) analyzed the differences in pass-through in a set of industries across a sample of European countries. He found that the direction of asymmetry varied across industries and countries. According to COUGHLIN and POLLARD (2004), the contrasting direction of the asymmetry highlights the importance of analyzing pass-through at the industry level. If the direction of asymmetry varies across industries then aggregation may obscure asymmetry that is present at the industry level.

Finally, the third potential source of nonlinearities is relative to *menu costs*.

- *Menu Costs*: because of the costs associated with changing prices, foreign exporters may leave their price in importer's currency unchanged if exchange rate changes are small. However, when exchange rate changes exceed some threshold i.e., with large magnitude, exporters do change their prices. Thus, according to menu costs hypothesis, ERPT may be asymmetric with respect to the size of the exchange rate shocks, since price adjustment is more frequent with large exchange rate changes than with small ones.

⁵Capacity constraints may also arise because of trade restrictions that limit imports, such as quotas or voluntary export restraints (see COUGHLIN and POLLARD (2004)).

This latter asymmetric dynamic behavior has been put forth empirically by COUGHLIN and POLLARD (2004). In their study on U.S. import prices of 30 industries, they found that most firms respond asymmetrically to large and small changes in the exchange rate with ERPT positively related to the size of the change. It is noteworthy that COUGHLIN and POLLARD (2004) use threshold dummy variables to distinguish between large and small exchange rate changes, in order to capture a possible asymmetric behavior in pass-through mechanism. The authors choose an arbitrary threshold value for 30 US industries to distinguish between small and large exchange rate changes. A large currency movement is defined as a change greater than 3%. In our paper, unlike COUGHLIN and POLLARD (2004), we propose to estimate a nonlinear smooth transition model where a grid search is used to select the appropriate threshold level instead of using an arbitrary threshold value.

Moreover, imported goods have to go through production or distribution processes before they reach consumers in domestic country. Thus, given the different pricing strategies along the distribution channel, this would affect substantially the transmission of exchange rate changes and, thereby, would account for asymmetric ERPT to consumer prices. As discussed in BACCHETTA and VAN WINCOOP (2002), the weakness of CPI inflation reaction to exchange rate changes is due, in part, to differences in the optimal pricing strategies of foreign producers and domestic wholesalers/retailers. Due to competitive pressure in the domestic market, domestic wholesalers import goods priced in foreign currency (PCP) and resell them in domestic currency (LCP). This would entail much lower ERPT to CPI inflation than expected. Also, substitution effect can occur following changes in relative prices. For example, if home currency is depreciating, domestic firms or wholesalers may reduce sourcing foreign products (since their price becomes higher), shifting towards substitute domestically produced goods. That way, consumer prices would be more insulated from exchange rate movements. Clearly then, the direction and the size of exchange changes would also affect pricing strategy of domestic wholesalers/retailers, and would account for possible asymmetric pass-through.

It is worth highlighting that the existing literature has focused notably on the asymmetries stemming from the size and the direction of exchange rate changes. Besides, there is some macroeconomic factors that would change foreign firms behavior,

and thus could be sources of nonlinearities in pass-through. One of these macroeconomic determinants is the inflation environment. As argued by TAYLOR (2000), the shift towards low and stable inflation regime has entailed a decline in the degree of ERPT in many industrialized countries. Accordingly, ERPT would be lower in a stable inflation environment than in higher inflation episodes. Therefore, one can think that dynamic behavior of pass-through depends upon inflation regime, which can be modeled in a non-linear way. To our knowledge, only three papers analyzed ERPT with respect to inflation level in a non-linear framework. Using a Phillips curve threshold model, PRZYSTUPA and WRÓBEL (2011) reject the hypothesis of an asymmetric pass-through related to inflation environment in Poland. On the other side, SHINTANI, TERADA-HAGIWARA, and TOMOYOSHI (2013) and NOGUEIRA JR. and LEON-LEDESMA (2008) found a strong positive correlation between ERPT and inflation in STR framework.⁶

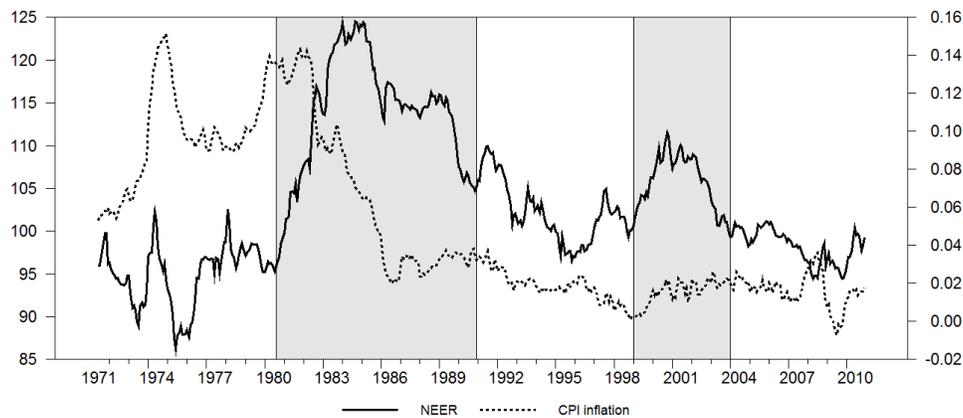
Another important source of pass-through nonlinearities is the business cycle. It is expected that when economy is booming, ERPT would be higher than in periods of slowdown. Intuitively, firms would find it easier to pass-through exchange rate changes when the economy is growing fast, rather than when it is in a recession and its sales are already falling. Empirically, this intuition was confirmed by GOLDFAJN and WERLANG (2000). Using a panel data model for 71 countries, they have found that depreciations have a higher pass-through to prices in periods of expansion. CORREA and MINELLA (2006) and PRZYSTUPA and WRÓBEL (2011) corroborated an asymmetric behavior between ERPT and growth in a Phillips curve threshold framework. Also, in their STR model, NOGUEIRA JR. and LEON-LEDESMA (2008) find that for three countries out of six pass-through responds nonlinearly to the output growth. Therefore, in this paper we aim to fill the gap in literature on sources of nonlinearities in ERPT. We analyze nonlinearities not only with respect to the size and the direction of exchange rate changes, but also to the inflation level and output growth.

Finally, it should be noted that dramatic exchange rate fluctuations, like those experienced by the European currencies in the 1980s, would provide an illustration of what seems to constitute an asymmetric pattern. The French Franc had depreciated by roughly 30 percent vis-a-vis major currencies (on a trade-weighted basis) between 1980-1985, followed by an appreciation of roughly the same magnitude by the end of 1990

⁶These two papers are discussed in section 4 with more details.

(see Figure 4.1). Similarly, at the launch of the EA, there has been a large depreciation of the Euro against the US dollar from 1999 till the last of 2001. After that date, the euro started appreciating to become a strong and well established currency. Thus, it is expected that these spectacular exchange rate developments may affect asymmetrically domestic prices, raising the question of the presence of a nonlinear dynamic in ERPT mechanism.⁷

Figure 4.1: Trade weighted nominal effective exchange rate and CPI inflation in France



Source: OCDE and IFS.

2.2. Analytical framework

Let us consider a foreign firm that exports its product i to an importing country. Under monopolistic competition, the first-order conditions for exporter profit maximization, with price, P_i , set in importing country currency, yield the following expression:

$$P_i = E \mu_i W_i^* \quad (4.1)$$

Where E is the exchange rate measured in units of the importer currency per unit of the foreign currency, μ_i is the markup of price over marginal cost W_i^* of foreign producer. The markup is defined as $\mu_i \equiv \eta_i / (1 - \eta_i)$, where η_i is the price elasticity of demand for

⁷Mussa (2005) argued that the large movement of the exchange rate of the US dollar against euro area currencies in 1980-1987 was significantly larger than the huge swing in the euro/dollar exchange rate since the beginning of 1999.

the good i in the importing country. As in BAILLIU and FUJII (2004), μ_i is assumed to depend essentially on demand pressures in the destination market: $\mu_i = \mu(Y)$, with Y is the income (expenditures) level in the importing country.

The log-linear form of equation (4.1) gives the standard ERPT regression traditionally tested throughout the exchange rate pass-through literature (see GOLDBERG and KNETTER (1997)):⁸

$$p_t = \alpha + \beta e_t + \psi y_t + \delta w_t^* + \varepsilon_t, \quad (4.2)$$

From equation (4.2), the ERPT coefficient is given by β and is expected to be bounded between 0 and 1. If $\beta = 1$, exporter markup will not respond to fluctuations of the exchange rate, price is set in foreign country currency (*producer-currency pricing*, PCP) and then the pass-through is complete. If $\beta = 0$, the ERPT is zero, since foreign firm decide not to vary the prices in the destination country currency and absorb the fluctuations within the markup. This is a purely *local-currency pricing* (LCP).

In the other hand, pricing strategies of firms depend not solely on demand conditions in the market. One can think that foreign firm may adjust price after exchange rate movements with respect to some macroeconomic factors. For instance, inflation environment, as argued by TAYLOR (2000), could influence the extent of ERPT. In a stable inflation environment ERPT would be lower than in higher inflation episodes. Thereby, a stable inflation environment in the destination country may lead exporters to adopt LCP strategy. Firms can accommodate currency changes within markup, leading to lesser extent of pass-through. While, when the importer experience high rates of inflation, exporter would change their pricing decision by adopting PCP strategy. Another important determinant of the ERPT mechanism is the business cycle. This latter might affect the transmission of exchange rate changes in a nonlinear way. In fact, firms are more willing to pass-through cost increases such as those coming from the exchange rate when the economy is growing faster, rather than when it is in a recession. Then, it is expected that ERPT would be higher in periods of prosperity than in periods of

⁸For simplicity, the good superscript i is dropped and time index t is added. Lower cases variables denote logarithms.

slowdown (see e.g. GOLDFAJN and WERLANG (2000)). Furthermore, foreign firm would adjust prices with respect to the magnitude or the direction of exchange rate movements. As mentioned above, exporters may leave their price unchanged if exchange rate changes are small due to the presence of menu costs. They change their prices only when the exchange rate change is above a given threshold. Thus, there will be differential effect of large versus small exchange rate changes on ERPT (see e.g. . Similarly, if firms attempt to keep competitiveness, faced with a depreciation of the importer currency, they tend to adjust markups to maintain market. Then an appreciation of the importing country's currency might cause higher pass-through than a depreciation.

Following these arguments, in our study, we assume that pricing strategy of foreign firms to depend on importer's macroeconomic environment in a nonlinear framework. We then consider $\kappa(M)$ as a function including those macroeconomic determinants such as inflation level, exchange rate direction or size and output growth. This macroeconomic dependence is seen as a firms' strategic decision on how much to translate exchange rate changes given different macroeconomic scenarios in the importing country. Taking into account these factors, we can re-write foreign firm markup as follow:

$$\mu_i = \mu(Y, E^{\kappa(M)}), \quad \kappa(M) \geq 0, \quad (4.3)$$

We can capture the arguments of equations (4.1) and (4.3) through a log-linear regression specification as follows:

$$\begin{aligned} p_t &= \alpha + \beta e_t + \psi y_t + \kappa(M) e_t + \delta w_t^* + \varepsilon_t \\ &= \alpha + [\beta + \kappa(M)] e_t + \psi y_t + \delta w_t^* + \varepsilon_t, \end{aligned} \quad (4.4)$$

According to the function $\kappa(M)$, there is an indirect channel of pass-through which depends on the macroeconomic environment. We have assumed macroeconomic factors affect firm's markup in a nonlinear way. We consequently consider that there is some threshold M^* which provide two extreme macroeconomic regimes. For example, if our macroeconomic variable is inflation rate, this enables us to distinguish between high and

low inflation environment regimes.

$$\kappa(M) = \begin{cases} 0 & \text{for } M \leq M^* \\ \phi & \text{for } M \geq M^* \end{cases} \quad (4.5)$$

According to (4.4) and (4.5), the degree of pass-through would be different and depends on whether the macroeconomic variable is above or below a threshold level. If the importing country has a small value (or a negative value if $M^* = 0$) for the macroeconomic variable, then ERPT is equal to β . If the importing country has a macroeconomic variable value above some threshold (or a positive value if $M^* = 0$), then ERPT is equal to $(\beta + \phi)$. We can see that ERPT is different depending on whether the macroeconomic determinant is above or below some threshold. For example, as mentioned in the literature, higher inflation environment would raise ERPT, however, with a stable inflation level pass-through would be lower. Thus, the advantage of equation (4.5) is to describe the changing behavior in the exchange rate in a nonlinear fashion, unlike previous empirical studies. Finally, it should be noted that the transition from one regime to the other is assumed to be smooth.

3. Empirical approach

3.1. Smooth transition regression models

To capture the nonlinearity in the exchange rate transmission, we use the family of smooth transition regression (STR) models as a tool. A STR model is defined as follows:

$$\begin{aligned} y_t &= \beta' \mathbf{z}_t + \phi' \mathbf{z}_t G(s_t; \gamma, c) + u_t \\ &= [\beta + \phi G(s_t; \gamma, c)]' \mathbf{z}_t + u_t, \end{aligned} \quad (4.6)$$

Where $u_t \sim \text{iid}(0, \sigma^2)$, $\mathbf{z}_t = (\mathbf{w}'_t, \mathbf{x}'_t)'$ is an $((m+1) \times 1)$ vector of explanatory variables with $\mathbf{w}'_t = (y_{t-1}, \dots, y_{t-d})'$ and $\mathbf{x}'_t = (x_{1t}, \dots, x_{kt})'$.⁹ $\beta = (\beta_0, \beta_1, \dots, \beta_m)'$ and $\phi = (\phi_0, \phi_1, \dots, \phi_m)'$ are the parameter vectors of the linear and the nonlinear part respectively. $G(s_t; \gamma, c)$ is the transition function bounded between 0 and 1, and depends upon the transition variable s_t , the slope parameter γ and the location parameter c .¹⁰ The transition variable s_t is an element of \mathbf{z}_t , and then is assumed to be a lagged endogenous variable ($s_t = y_{t-d}$) or an exogenous variable ($s_t = x_{kt}$). We note that the equation (4.6) can be interpreted also as a linear model with stochastic time-varying coefficients $\beta + \phi G(s_t; \gamma, c)$ depending on the value of s_t .

There are two standard choice of the transition function:

- Logistic Function

$$G(s_t; \gamma, c) = [1 + \exp\{-\gamma(s_t - c)\}]^{-1} \quad (4.7)$$

- Exponential Function

$$G(s_t; \gamma, c) = 1 - \exp\{-\gamma(s_t - c)^2\} \quad (4.8)$$

Equations (1) and (2) jointly define the logistic STR (LSTR) model and the pattern formed jointly by (1) and (3) is called the exponential STR (ESTR) model. In Both models, the parameter c can be interpreted as the threshold between two extremes regimes corresponding to $G(s_t; \gamma, c) = 0$ and $G(s_t; \gamma, c) = 1$. For the LSTR model, the nonlinear coefficients would take different values depending on whether the transition

⁹When \mathbf{x}_t is absent from (1) and $s_t = y_{t-d}$, the STR model becomes a univariate smooth transition autoregressive (STAR) model.

¹⁰The parameter γ is also called the speed of transition which determines the smoothness of the switching from one regime to the other.

variable is below or above the threshold. So, the parameters $\phi + \theta G(s_t; \gamma, c)$ changes monotonically as a function of s_t from β to $(\beta + \phi)$. In this sense, as $(s_t - c) \rightarrow -\infty$, $G(s_t; \gamma, c) \rightarrow 0$ and coefficients correspond to β ; if $(s_t - c) \rightarrow +\infty$, $G(s_t; \gamma, c) \rightarrow 1$ and coefficients are equal to $(\beta + \phi)$; and if $s_t = c$, then $G(s_t; \gamma, c) = 1/2$ and coefficients will be $(\beta + \phi/2)$. It should be noted that LSTR model would follow the same pattern as the threshold model described in the theoretical model (equation (4.5)) but assuming a smooth adjustment between across regimes. One feature of LSTR model is that when $\gamma \rightarrow \infty$, LSTR model approaches the two-regime switching regression model with an abrupt transition (Threshold Regression). But when $\gamma = 0$, the transition function $G(s_t; \gamma, c) \equiv 0$, and thus the LSTR model reduces to a linear model.

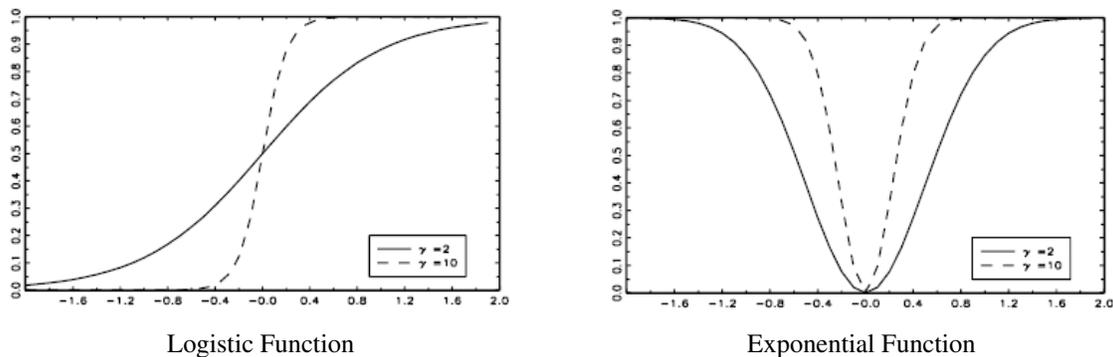
Concerning ESTR model, this specification is appropriate in situations in which the dynamic behavior is different for large and small values of s_t - what matters is the magnitude of shock, if they are large or small. In other words, the coefficient changes depending on whether s_t is near or far away from the threshold, regardless of whether this difference $(s_t - c)$ is positive or negative. Therefore, the exponential transition function $G(s_t; \gamma, c) \rightarrow 1$ as $(s_t - c) \rightarrow \pm\infty$ and the coefficients of the model will be equal $(\beta + \phi)$. And if $s_t = c$, $G(s_t; \gamma, c) = 0$ and coefficients becomes β . A drawback of ESTR specification is that for either $\gamma \rightarrow \infty$ and $\gamma \rightarrow 0$, the model becomes practically linear and thus it does not nest a threshold regression model (with steep transition) as a special case.

The implied nonlinear dynamics under logistic and exponential functions are drastically different Figure 4.2. LSTR model is pertinent in describing asymmetric dynamic behavior. As mentioned in the STR literature (VAN DIJK, TERÄSVIRTA, and FRANSES (2002)), when modeling business cycle, LSTR can describe processes whose dynamic properties are different in expansions from what they are in recessions. For example, if the transition variable s_t is a business cycle indicator (such as output growth), and if $c \simeq 0$, the model distinguishes between periods of positive and negative growth, that is, between expansions and contractions. On the other hand, an ESTR allow for symmetric dynamics with respect to negative or positive deviations of s_t from the threshold level. The function rather depends on whether the transition variable is close or far away from the threshold c . Exponential specification was popularly employed in analyzing the nonlinear adjustment of real exchange rates (see MICHAEL, NOBAY,

and PEEL (1997), TAYLOR and PEEL (2000), TAYLOR, PEEL, and SARNO (2001), and KILIAN and TAYLOR (2003), among others).

Therefore, we must be careful in our implementation of these models in our ERPT analysis. LSTR and ESTR models must allow respectively for asymmetric and symmetric response of domestic prices to exchange rate changes with respect to negative and positive deviations of s_t from c . For example, when considering exchange rate as transition variable and $c \simeq 0$, LSTR model can account for asymmetric ERPT during currency appreciations and depreciations episodes. For ESTR model interpretation is different, and what matters is the size of exchange rate change. According to ERPT literature, firms are willing to absorb small changes in exchange rate rather than larger ones due to the presence of "menu costs". Thus, the costs of changing prices may result in asymmetric pass-through for large and small exchange rate shocks (COUGHLIN and POLLARD (2004)). In such case, ESTR specification would be more appropriate in describing this kind of non-linearity (NOGUEIRA JR. and LEON-LEDESMA (2008)). We will see later that our choice for relevant transition function must be also conducted with non-linearity specification tests in addition to the economic intuition.

Figure 4.2: Transition Functions



3.2. Modelling strategy of STR models

The modeling procedure follows TERÄSVIRTA (1994) approach and is similar to the modeling cycle for linear models of Box and Jenkins (1970). It is consisting of three stages: specification, estimation, and evaluation:

3.2.1. Specification stage

As a starting point for the analysis, adequate linear representation must be specified. This can be modelled by using a VAR framework. For lag selection, we adopt a general-to-specific approach, as suggested by VAN DIJK, TERÄSVIRTA, and FRANSES (2002), to reach the final specification. We start with a model with maximum lag length $N = 4$, and sequentially we remove the lagged variables for which the t -statistic of the corresponding parameter is less than 1.0 in absolute value. Second step of specification consists in testing for non-linearity, choosing the appropriate s_t and the most suitable form of the transition function, i.e. LSTR or ESTR models.

Linearity is tested against a STR model with a predetermined transition variable. Economic theory may give an idea of which variables should be selected as s_t . Alternatively, the test is repeated for each variable in the set of potential transition variables, which is usually a subset of the elements in \mathbf{z}_t . If the null hypothesis of linearity is rejected for at least one of the candidate models, the model against which the rejection is strongest is chosen to be the STR model to be estimated. Once linearity has been rejected and a transition variable subsequently selected, the final decision to be made at this stage concerns the appropriate form of the transition function.

In order to derive a linearity test, TERÄSVIRTA (1994, 1998) suggest to approximate the logistic function (4.7) in (4.6) by a third-order Taylor expansion around the null hypothesis $\gamma = 0$. The resulting test has power against both the LSTR and ESTR models. Assuming that the transition variable s_t is an element in \mathbf{z}_t and let $\mathbf{z}_t = (1, \tilde{\mathbf{z}}_t)'$, where $\tilde{\mathbf{z}}_t'$ is an $(m \times 1)$.

Taylor approximation yields the following auxiliary regression:

$$y_t = \alpha_0' \mathbf{z}_t + \sum_{j=1}^3 \alpha_j' \tilde{\mathbf{z}}_t s_t^j + u_t^*, \quad t = 1, \dots, T, \quad (4.9)$$

Where $u_t^* = u_t + R_3(\gamma, c, s_t) \theta' \mathbf{z}_t$, with $R_3(\gamma, c, s_t)$ the residual of Taylor expansion. The null hypothesis of linearity is $H_0 : \alpha_1 = \alpha_2 = \alpha_3 = 0$. LUUKKONEN, SAIKKONEN,

and TERÄSVIRTA (1988) suggest a Lagrange Multiplier (LM) statistic with a standard asymptotic $\chi^2(3m)$ distribution under the null hypothesis. In small and moderate samples, the χ^2 -statistic may be heavily oversized. The F version of the test is recommended instead, which has an approximate F -distribution with $3m$ and $T - 4m - 1$ degrees of freedom under H_0 (VAN DIJK, TERÄSVIRTA, and FRANSES (2002)). Linearity tests are executed for each of the candidate potential transition variables. If the null hypothesis is rejected for several transition variables, select the one with the strongest test rejection (the smallest p -value). The logic behind this suggestion is that the rejection of H_0 is stronger against the correct alternative than other alternative models. However, if several small p -values are close to each other, it may be useful to proceed by estimating the corresponding STR models and leaving the choice between them to the evaluation stage.

Once linearity has been rejected, one has to choose whether an LSTR or an ESTR model should be specified. The choice between these two types of models can be based on the auxiliary regression (equation (4.9)). TERÄSVIRTA (1994, 1998) suggested that this choice can be based on testing the following sequence of nested null hypotheses:

1. Test $H_{04} : \alpha_3 = 0$
2. Test $H_{03} : \alpha_2 = 0 | \alpha_3 = 0$
3. Test $H_{02} : \alpha_1 = 0 | \alpha_2 = \alpha_3 = 0$

According to TERÄSVIRTA (1994), the decision rule is the following: if the test of H_{03} yields the strongest rejection measured in the p -value, choose the ESTR model. Otherwise, select the LSTR model. All three hypotheses can simultaneously be rejected at a conventional significance level, that is why the strongest rejection counts. This procedure was simulated in TERÄSVIRTA (1994) and appeared to work satisfactorily. According to VAN DIJK, TERÄSVIRTA, and FRANSES (2002), recent increases in computational power have made these decision rules less important in practice. Since it is easy to estimate a number of both LSTAR and ESTAR models and to choose between them at the evaluation stage by misspecification tests. In practice, this is a sensible way of proceeding if the test sequence does not provide a clear-cut choice between the two

alternatives in the sense that p-values of the test of H_{03} , on the one hand, and of H_{02} or H_{04} on the other, are close to each other. Nevertheless, carrying out the tests still be recommended even if the actual decision were postponed to the evaluation stage of the modelling strategy.

3.2.2. Estimation stage

The parameters of the STR model are estimated by non-linear least squares (NLS) estimation technique which provides estimators that are consistent and asymptotically normal. As discussed in VAN DIJK, TERÄSVIRTA, and FRANSES (2002), under the assumption that the errors are normally distributed, NLS is equivalent to maximum likelihood. Otherwise, the NLS estimates can be interpreted as quasi maximum likelihood estimates. Finding good starting values is crucial in this procedure. Thus, STR literature suggests to construct a grid search for estimating γ and c . The values for the grid search for γ were set between 0 and 100 for increments of 1, whereas c was estimated for all the ranked values of the transition variable s_t . For each value of γ and c the residual sum of squares is computed. The values that correspond to the minimum of that sum are taken as starting values into the NLS procedure. This procedure increases the precision of the estimates and ensures faster convergence of the NLS algorithm. It should also be noted that when constructing the grid, γ is not a scale-free. The transition parameter γ is therefore standardized by dividing it by the sample standard deviation of the transition variable s_t , which will call $\hat{\sigma}_s$. Then, the transition functions become:

$$G(s_t; \gamma, c) = \begin{cases} \left[1 + \exp(-(\gamma/\hat{\sigma}_s)(s_t - c)) \right]^{-1} & \text{for Logistic Function} \\ 1 - \exp(-(\gamma/\hat{\sigma}_s)(s_t - c)^2) & \text{for Exponential Function} \end{cases} \quad (4.10)$$

3.2.3. Evaluation stage

In the final stage, the quality of the estimated STR model should check against misspecification as in the case of linear models. Several misspecification tests are used in the STR literature, such as LM test of no error autocorrelation, LM-type test of no

ARCH and Jarque-Bera normality test. EITRHEIM and TERÄSVIRTA (1996) suggested two additional LM-type misspecification tests: an LM test of no remaining nonlinearity and LM-type test of parameter constancy. We briefly describe the two latter tests.

Test of no remaining nonlinearity: After estimating STR model parameters, it is important to ask whether some nonlinearity remains unmodeled. The test assumes that the type of the remaining nonlinearity is again of the STR type. The alternative can be defined as:

$$y_t = \beta' \mathbf{z}_t + \phi' \mathbf{z}_t G(s_{1t}; \gamma_1, c_1) + \psi' \mathbf{z}_t H(s_{2t}; \gamma_2, c_2) + u_t, \quad (4.11)$$

where H is another transition function and $u_t \sim \text{iid}(0, \sigma^2)$. To test this alternative, the following auxiliary model is used:

$$y_t = \alpha_0' \mathbf{z}_t + \phi' \mathbf{z}_t G(s_{1t}; \gamma_1, c_1) + \sum_{j=1}^3 \alpha_j' \tilde{\mathbf{z}}_t s_{2t}^j + u_t^*, \quad (4.12)$$

The null hypothesis of no remaining nonlinearity is that $\alpha_1 = \alpha_2 = \alpha_3 = 0$. The choice of s_{2t} can be a subset of available variables in \mathbf{z}_t or it can be s_{1t} . It is also possible to exclude certain variables from the second nonlinear part by restricting the corresponding parameter to zero. The resulting F -statistics are given in the same way as for the test on linearity.

Test of parameter constancy: This is a test against the null hypothesis of constant parameters against the alternative of smooth continuous change in parameters.¹¹ To

¹¹This is different from parameter constancy test in linear model, where the alternative is a single structural break. The present alternative does, however, contain a structural break as a special case.

consider the test, rewrite (1) as follows:

$$y_t = \beta(t)' \mathbf{z}_t + \phi(t)' \mathbf{z}_t G(s_{1t}; \gamma_1, c_1) + u_t, \quad (4.13)$$

where

$$\beta(t)' = \beta + \lambda_\beta + H_\phi(\gamma_\beta, c_\beta, t^*) \quad (4.14)$$

and

$$\phi(t)' = \phi + \lambda_\phi + H_\phi(\gamma_\phi, c_\phi, t^*) \quad (4.15)$$

with $t^* = t/T$ and $u_t \sim \text{iid}(0, \sigma^2)$. $H_\beta(\gamma_\beta, c_\beta, t^*)$ and $H_\phi(\gamma_\phi, c_\phi, t^*)$ are transition functions with $s_t = t^*$. The null hypothesis of no change in parameters is $\gamma_\beta = \gamma_\phi = 0$. The parameters γ and c are assumed to be constant. The following nonlinear auxiliary regression is used:

$$y_t = \alpha'_0 \mathbf{z}_t + \sum_{j=1}^3 \alpha'_j \tilde{\mathbf{z}}_t(t^*)^j + \sum_{j=1}^3 \alpha'_{j+3} \tilde{\mathbf{z}}_t(t^*)^j G(s_t; \gamma, c) + u_t^*, \quad (4.16)$$

where $\alpha_j = 0$, $j = 1, \dots, 6$, if and only if the null hypothesis $\gamma_\beta = \gamma_\phi = 0$ holds. As usual, the F -version of the LM test is recommended instead of the χ^2 variants which may be heavily oversized in small samples.

In the STR literature, error autocorrelation, parameter nonconstancy and remaining nonlinearity tests are the most obvious ones used in the evaluation stage,

nevertheless, other tests such as the LM-type test for the null hypothesis of no ARCH and the Jarque-Bera normality test may be useful.¹²

4. Empirical literature of STR pass-through model

The empirical literature that utilizing the STR models in examining the extent of ERPT is to date relatively scarce, although it constitutes an important extension to look for as discussed by HERZBERG, KAPETANIOS, and PRICE (2003). Only a very few number of studies have tested for nonlinearities and asymmetries in this context. Essentially, we can mention three studies who estimates a STR pass-through model: SHINTANI, TERADA-HAGIWARA, and TOMOYOSHI (2013) for US, NOGUEIRA JR. and LEON-LEDESMA (2008, 2011) for six countries adopting Inflation Target (IT) regime and HERZBERG, KAPETANIOS, and PRICE (2003) for UK. The latter papers were interested in measuring the ERPT into import prices but did not find any evidence of nonlinearity. Therefore, we will introduce only the first two studies in this section, namely, SHINTANI, TERADA-HAGIWARA, and TOMOYOSHI (2013) and NOGUEIRA JR. and LEON-LEDESMA (2008, 2011).

4.1. SHINTANI, TERADA-HAGIWARA, and TOMOYOSHI (2013)

In their paper, the authors use a STAR model to measure US domestic price adjustment to exchange rate movements from 1975 to 2007 using monthly data. Primarily, they estimate a bivariate version of a STAR model specified as follow:

$$\begin{aligned} \pi_t = \phi_0 + \sum_{j=1}^N \phi_{1,j} \pi_{t-j} + \sum_{j=0}^{N-1} \phi_{2,j} \Delta(e_{t-j} + p_{t-j}^*) \\ + \left(\sum_{j=1}^N \phi_{3,j} \pi_{t-j} + \sum_{j=0}^{N-1} \phi_{4,j} \Delta(e_{t-j} + p_{t-j}^*) \right) G(s_t; \gamma) + \varepsilon_t, \end{aligned} \quad (4.17)$$

¹²Jarque-Bera normality test is sensitive to outliers, and the result should be considered jointly with a visual examination of the residuals.

Where π_t is the inflation rate of the producer price index and $\Delta(e_t + p_t^*)$ US dollar prices paid by the US importer.¹³ According to their theoretical model the ERPT is a symmetric function of the past inflation rates around zero. To capture this feature, an exponential *U-shaped* symmetric transition function is used:

$$G(s_t; \gamma) = 1 - \exp\{-\gamma s_t^2\}, \quad (4.18)$$

Only one transition variable is used in this empirical analysis, which is a moving average of the past inflation rates, $s_t = d^{-1} \sum_{j=1}^d \pi_{t-j}$. In addition to the ESTAR model, they also consider another STAR model constructed from a combination of two logistic functions, which gives a different *U-shaped* transition function. Thus, the transition function in this is given by:

$$G(s_t; \gamma_1, \gamma_2, c) = (1 + \exp\{-\gamma_1(s_t - c_1)\})^{-1} + (1 + \exp\{-\gamma_2(s_t + c_2)\})^{-1} \quad (4.19)$$

SHINTANI, TERADA-HAGIWARA, and TOMOYOSHI (2013) call it dual LSTAR (DLSTAR) model to emphasize the presence of two logistic functions, which is different from the STAR model with "second-order" logistic function¹⁴: $G(s_t; \gamma, c_1, c_2) = (1 + \exp\{-\gamma(s_t - c_1)(s_t - c_2)\})^{-1}$. According to SHINTANI, TERADA-HAGIWARA, and TOMOYOSHI (2013), there are two reasons behind the use of DLSTR model (equation (4.19)). First, as mentioned above, the transition function in the ESTAR model collapses to a constant when γ approaches infinity, and then the model does not nest the TAR (Threshold Autoregressive) model with an abrupt transition. In contrast, the DLSTAR model includes the TAR model by letting γ_1, γ_2 tend to infinity. Second, and more importantly, the model can incorporate both symmetric ($\gamma_1 = \gamma_2$ and $c_1 = c_2$) and asymmetric ($\gamma_1 \neq \gamma_2$ and $c_1 \neq c_2$) adjustments between the positive and negative regions.

¹³SHINTANI, TERADA-HAGIWARA, and TOMOYOSHI (2013) employ the producer price index rather than the consumer price index. In their model, they consider that the domestic price is the price at which the final good producer sells its product.

¹⁴See VAN DIJK, TERÄSVIRTA, and FRANSES (2002).

Therefore, this enables investigating a symmetric relationship between the ERPT and the inflation rate.

Concerning their results, the authors found that the degree of ERPT becomes largest when the transition variable becomes above 2% in absolute term. They detect three distinct high ERPT episodes. The first period in which ERPT was high is during the second oil shock episodes. During the 1980s and 1990s, there was a relatively low pass-through except for the early 1990s when the producer price index was relatively volatile. The last period corresponds to the beginning of 2000, the ERPT becomes high again due to the increased volatility of inflation. Therefore, SHINTANI, TERADA-HAGIWARA, and TOMOYOSHI (2013) conclude that the period of low ERPT is likely to be associated with the low inflation environment, and vice versa.

4.2. NOGUEIRA JR. and LEON-LEDESMA (2008, 2011)

NOGUEIRA JR. and LEON-LEDESMA (2008, 2011) investigate the ERPT into CPI inflation for a set of emerging and developed IT countries. Using monthly data, the period of estimation is ranging from 1983 to 2005 for the developed economies, and 1992 to 2005 for the emerging markets. Empirically, the authors consider the following model:

$$\begin{aligned} \pi_t = & \beta_0 + \sum_{j=1}^N \beta_{1,j} \pi_{t-j} + \sum_{j=0}^N \beta_{2,j} \Delta p_{t-j}^{imp} + \sum_{j=0}^N \beta_{3,j} \Delta y_{t-j} \\ & + \sum_{j=0}^N \beta_{4,j} \Delta e_{t-j} + \left(\beta_0^* + \sum_{j=0}^N \beta_{4,j}^* \Delta e_{t-j} \right) G(s_t; \gamma, c) + \varepsilon_t, \end{aligned} \quad (4.20)$$

Where π_t is the consumer prices index (CPI) inflation rate, Δp_t^{imp} is the change in import prices (in foreign currency), Δy_t is the output growth and Δe_t is the rate of exchange rate depreciation. NOGUEIRA JR. and LEON-LEDESMA (2008, 2011) experiment both traditional logistic and exponential (equations (4.7) and (4.8) respectively) as a transition function. They test several potentially important transition variables in order to capture possible nonlinearities and asymmetries in ERPT. So, they consider a set of macroeconomic variables which affecting the ERPT mechanism,

namely inflation rate, exchange rate, output growth and two measures of macroeconomic instability.¹⁵ For more accuracy, an ESTR model was used to capture asymmetry of pass-through with respect to the size of exchange rate change, i.e. asymmetry between large and small shocks. On the other hand, LSTR model was employed for the rest of transition variables (inflation, output and instability measures) as dynamic behavior would be different on each side of the threshold.

NOGUEIRA JR. and LEON-LEDESMA (2008, 2011) results highlight several sources of linearities in ERPT which vary considerably across countries. First, 4 out of 6 countries show a positive relationship between ERPT and the inflation level. According to the authors, the adoption of IT, which was followed by lower inflation in this countries sample, has contributed in moderating pass-through. Also, the authors find that ERPT seems to increase in periods of confidence crisis, which highlights the importance of a stable macroeconomic environment in reducing ERPT. When considering exchange rate as transition variable only two countries indicate a positive correlation between pass-through and exchange rate change magnitude. Finally, ERPT seems to be affected nonlinearly by output growth, when the economy is growing above some threshold, ERPT would be higher. This latter result is valid for 3 out of 6 countries in the sample.

5. Empirical specification and data

In our empirical analysis, we define a *STR pass-through equation* that enables us to test the presence of a nonlinear ERPT mechanism in the EA countries. As a matter of fact, the theoretical model (4.4) is designed to test the ERPT to import prices, i.e. the so-called “first step pass-through”, while our paper instead deals with the responsiveness of consumer prices. Thus, as recommended by BAILLIU and FUJII (2004), typical pass-through equations (such as equation (4.4)) could be adjusted in order to have all the elements of a backward-looking Phillips curve. Mainly, there are two issues which we consider here: first, the inertial behavior of inflation. This could be accomplished by including lags of inflation (π_{t-j}) as explanatory variables in

¹⁵NOGUEIRA JR. and LEON-LEDESMA (2008, 2011) use two potential indicators of macroeconomic instability: real interest rates differentials to the United States and Emerging Markets Bond Index Plus (EMBI+) spreads which is a leading indicator of confidence crises.

the empirical specification (backward-looking inflation). Second, a proxy for changes in domestic demand conditions should enter the pass-through equation. We use the changes in real GDP (Δy_t) to capture this effect.¹⁶ Once these two elements have been considered, our STR pass-through equation can be described as a nonlinear backward-looking Phillips curve as follows:

$$\begin{aligned} \pi_t = & \alpha + \sum_{j=1}^N \lambda_j \pi_{t-j} + \sum_{j=0}^N \psi_j \Delta y_{t-j} + \sum_{j=0}^N \delta_j \Delta w_{t-j}^* \\ & + \sum_{j=0}^N \beta_j \Delta e_{t-j} + \left(\sum_{j=0}^N \phi_j \Delta e_{t-j} \right) G(s_t; \gamma, c) + \varepsilon_t, \end{aligned} \quad (4.21)$$

Where π_t is the CPI inflation rate, Δw_t^* is the changes in foreign producer cost, Δy_t is the output growth and Δe_t is the rate of depreciation of the nominal effective exchange rate. $G(s_t; \gamma, c)$ is the transition function driving the nonlinear dynamic. According to (4.21), we can define both short- and long-run time-varying ERPT coefficients.

- Short-run pass-through:

$$SR\ ERPT = \beta_0 + \phi_0 G(s_t; \gamma, c) \quad (4.22)$$

- Long-run pass-through:

$$LR\ ERPT = \frac{\sum_{j=0}^N \beta_j + \sum_{j=0}^N \phi_j G(s_t; \gamma, c)}{1 - \sum_{j=1}^N \lambda_j} \quad (4.23)$$

¹⁶Also, we can use the output gap computed as the difference between actual and potential output (constructed with a Hodrick-Prescott filter) instead of real output growth. This does not alter the estimates of pass-through.

$G(s_t; \gamma, c)$ is assumed to be either logistic or exponential function as specified in the equations (4.7) and (4.8). For the LSTR model, ERPT coefficient would take different values depending on whether the transition variable is below or above the threshold.

- If $(s_t - c) \rightarrow -\infty$, pass-through elasticities are equal to:

$$SR\ ERPT = \beta_0 \tag{4.24}$$

and

$$LR\ ERPT = \frac{\sum_{j=0}^N \beta_j}{1 - \sum_{j=1}^N \lambda_j} \tag{4.25}$$

- If $(s_t - c) \rightarrow +\infty$, pass-through coefficients become:

$$SR\ ERPT = \beta_0 + \phi_0 \tag{4.26}$$

and

$$LR\ ERPT = \frac{\sum_{j=0}^N \beta_j + \sum_{j=0}^N \phi_j}{1 - \sum_{j=1}^N \lambda_j \pi_{t-j}} \tag{4.27}$$

In the case of the ESTR model, pass-through elasticities change depending on whether s_t is near or far away from the threshold c , regardless of whether the difference $(s_t - c)$ is positive or negative. Therefore, if $(s_t - c) \rightarrow \pm\infty$, short-run and long-run ERPT correspond, respectively, to equation (4.26) and (4.27); and if $s_t = c$, short-run and long-run pass-through coefficients will be equal to (4.24) and (4.25) respectively.

The STR pass-through equation (4.21) is estimated for 12 euro area countries (Austria, Belgium, Germany, Spain, Finland, France, Greece, Ireland, Italy, Luxembourg, Netherlands and Portugal), using quarterly data spanning the period 1975:1 to 2010:4. All the data we use are taken from the OECD's *Economic Outlook* database, except for exchange rate series which are obtained from *International Financial Statistics* (IFS) of the International Monetary Fund (IMF). Inflation rates series represents the change in consumer prices index (CPI). Output growth is constructed using the rate of growth of the real GDP. The nominal exchange rate is defined as domestic currency units per unit of foreign currencies, which implies that an increase represents a depreciation for home country. Finally, to capture changes in foreign costs, we construct a typical export partners cost proxy (W_t^*) that used throughout the ERPT literature (see BAILLIU and FUJII (2004) and CAMPA and GOLDBERG (2005)): $W_t^* = Q_t \times W_t / E_t$, where Q_t is the unit labor cost based real effective exchange rate, W_{it} is the domestic unit labor cost and E_t is the nominal effective exchange rate. Taking the logarithm we obtain the following expression: $w_{it}^* \equiv q_t + w_t - e_t$. Since the nominal and real effective exchange rate series are trade weighted, we obtain a measure of foreign firms' costs with each partner weighted by its importance in the domestic country's trade.

We have checked the possibility of cointegrating relationship among our variables in ERPT equation (4.4). Individual series in level are non-stationary but do not appear to be cointegrated according to ENGLE and GRANGER (1987) tests (henceforth EG test) results. According to DE BANDT, BANERJEE, and KOZLUK (2007), the long run equilibrium relation may be restored once we take into account the possibility of structural breaks in the data. Since we use long sample period (144 time observations for each country), we employ GREGORY and HANSEN (1996), henceforth GH test, methodology which test the null of no cointegration against the alternative of cointe-

gration with an estimated structural break.¹⁷ In spite of allowing for possible breaks in ERPT equation, we failed to reject the hypothesis of no cointegration for most of country sample (see results in Table D.5 in Appendix D.1). As a result, log differences of the variables are used in the estimation the STR pass-through equation as shown in equation (4.21). Augmented Dickey Fuller (ADF) tests suggest that variables in differences are appropriately described as stationary series. In addition to ADF tests, we have implemented ZIVOT and ANDREWS (1992) and LUMSDAINE and PAPELL (1997) unit root tests which allow for possible breaks in series (see Tables D.1, D.2, D.3 and D.4 in Appendix D.1).¹⁸

To determine the lag length of the variables, we follow VAN DIJK, TERÄSVIRTA, and FRANSES (2002) by adopting a general-to-specific approach to select the final specification. Then, we start with a model with maximum lag length of $N = 4$, and then dropping sequentially the lagged variables for which the t -statistic of the corresponding parameter is less than 1.0 in absolute value. The next step consists in testing for non-linearity, selecting the appropriate s_t and choosing the adequate form of the transition function, namely logistic or exponential. In our empirical analysis, three potential transition variables are considered: inflation rate, exchange rate and output growth. Then, the linearity tests are conducted for the respective delayed variables, i.e. π_{t-i} , Δe_{t-i} and Δy_{t-i} for lag length up to four periods ($i = 4$).

As mentioned in the specification stage (section 3.2.1), we follow TERÄSVIRTA (1994, 1998) procedure. Tables D.7, D.8 and D.9 provides the p -values of the F version of the LM test with the different lags for the transition variables. In the first row, we report the test of the null hypothesis of linearity against the alternative of STR non-linear model.¹⁹ The following rows in each table show the sequence of null hypotheses for choosing the LSTR or the ESTR model. The decision rule for the test is as follow: if the p -value of the test corresponding to H_{03} is the smallest, we choose an ESTR model, while in all other cases an LSTAR model should be selected. According to TERÄSVIRTA

¹⁷Two alternative versions of GH test are used: a first model which allows for break in constant and a second model with break both in constant and slope.

¹⁸LUMSDAINE and PAPELL (1997) test is the extension of ZIVOT and ANDREWS (1992) model by allowing for two structural breaks under the alternative hypothesis (instead of a single break).

¹⁹Additionally, in our choice of the transition variable we also test whether some nonlinearity remains unmodelled with the test of no additive nonlinearity at the evaluation stage.

(1994), all three hypotheses (H_{04} , H_{03} and H_{02}) can simultaneously be rejected at a conventional significance level, that is why the strongest rejection counts.

According to the STR literature, it is also recommended to estimate a number of both LSTAR and ESTAR models and to choose between them at the evaluation stage by misspecification tests, such as error autocorrelation, parameter nonconstancy and remaining nonlinearity. This way of proceeding is advocated if the test sequence does not provide a clear-cut choice between the two alternatives, that is, p -values of the test of H_{03} , on the one hand, and of H_{02} or H_{04} on the other, are close to each other. Therefore, the final decision can be postponed to the evaluation stage of the modeling strategy as recommended by TERÄSVIRTA (1994, 1998, 2004) and VAN DIJK, TERÄSVIRTA, and FRANSES (2002).

It is noteworthy that our choice for relevant transition function must be also conducted with the economic intuition in addition to non-linearity specification tests. As mentioned in the pass-through literature, exchange rate transmission would be lower in a stable inflation environment than in a higher inflation periods, which is a proof of regime-dependence of ERPT to inflation environment. Then, when considering lagged inflation as transition variable, LSTR model would be more appropriate in describing this asymmetric behavior. Similarly, when considering output growth as transition variable, the LSTR specification is preferred since pass-through mechanisms could be different in expansions from what they are in recessions. Finally, for the exchange rate case, there are two types of linearities that must be modeled. On one hand, we use a LSTR model to can account for ERPT asymmetry during currency appreciations and depreciations episodes. On the other hand, an ESTR is chosen to capture non-linearity in pass-through with respect to large and small exchange rate fluctuations.

6. Main Empirical Results

6.1. Linear model results

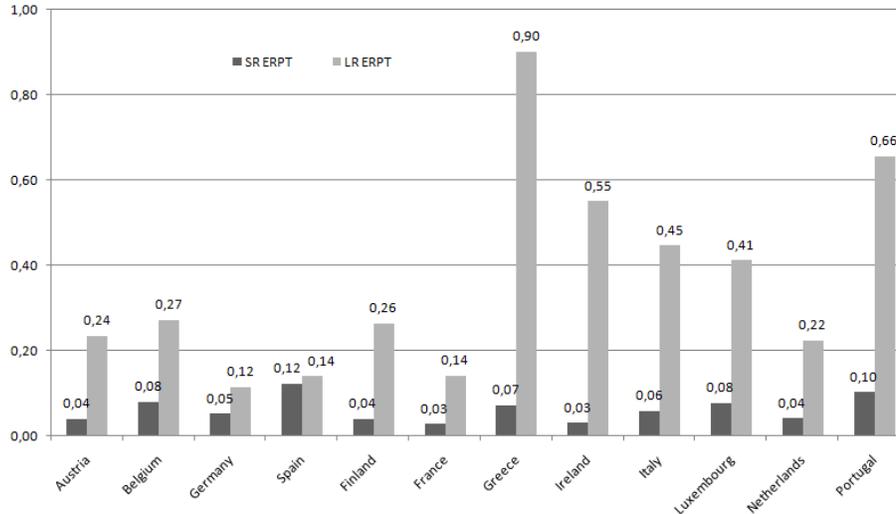
We begin our analysis by estimating a linear version of ERPT model which corresponds to the equation (4.21) without the non-linear part. The objective is twofold: first, we measure the extent of pass-through and compare this with results from the existing

literature. Second, this enable us later to make a comparison with the nonlinear STR model from a statistical point of view, such as a comparison of R^2 , sum of squared residuals and Akaike Information Criterion (AIC). Therefore, we estimate the following linear ERPT equation:

$$\pi_t = \alpha + \sum_{j=1}^4 \lambda_j \pi_{t-j} + \sum_{j=0}^4 \psi_j \Delta y_{t-j} + \sum_{j=0}^4 \delta_j \Delta w_{t-j}^* + \sum_{j=0}^4 \beta_j \Delta e_{t-j} + \varepsilon_t, \quad (4.28)$$

Figure 4.3 reports OLS estimates of the short- and long-run ERPT for the 12 EA countries (detailed estimation results are presented in the Table D.6 in the Appendix D.2). Our results suggest a moderate effect of exchange rate changes on consumer price inflation in the short run. The average of short-run elasticity in the EA sample is 0.06, suggesting that 1% increase in the rate of currency depreciation leads to 0.06% increase in the inflation rate. The higher rate was recorded in Spain with 0.12%, and the lower was found in France and Ireland with 0.03%. Our results are in line with estimates in the literature of pass-through. Using dynamic panel data model, BAILLIU and FUJII (2004) found a pass-through to CPI inflation equal to 0.08% . This lower rate is valid for 11 industrialized countries among them there are six euro area countries, namely Belgium, Finland, France, Italy, Netherlands and Spain. In a large database including 1979-2000 quarterly data for 71 countries, CHOUDHRI and HAKURA (2006) provide evidence of low short run ERPT for low inflation countries such as EA sample.²⁰ We can notice that, in our study, we have found nearly the same elasticity as in CHOUDHRI and HAKURA (2006), especially for Austria, Belgium and Germany (respectively 0.04, 0.08 and 0.05 per cent).

²⁰CHOUDHRI and HAKURA (2006) classify their sample of countries into three groups: low inflation, moderate inflation and high inflation. Low, moderate and high inflation groups are defined as consisting of countries with average inflation rates less than 10%, between 10 and 30% and more than 30%, respectively. All the 12 EA countries were included in the low inflation group except Greece and Portugal.

Figure 4.3: Estimated short-run and long-run ERPT from linear model

Sources: Personal calculation.

In the long run, as expected, ERPT is higher than in the short run due to the gradual adjustment of prices to exchange rate movements. We note that the rates of pass-through vary substantially across countries, ranging from 0.12 in Germany to 0.90 in Greece. According to Figure 4.3, five countries out of twelve have price reaction exceeding 0.40%. In their sample of 20 industrialized countries, GAGNON and IHRIG (2004) found that Greece and Portugal have the highest degree of pass-through with 0.43 and 0.52, respectively, which corroborates our estimation results. In CHOUDHRI and HAKURA (2006), these two countries has been classified among countries with medium inflation rate (between 10 and 30%), and, consequently, the long-run pass-through was found to be higher compared to low inflation countries.²¹ As a result, the average long-run rate of pass-through in our 12 EA countries - which is equal to 0.36% - is found to be close to the average of pass-through elasticity in medium inflation countries in CHOUDHRI and HAKURA (2006) - which is equal to 0.35. Obviously, this is due to the higher rate of ERPT in Greece and Portugal among our EA countries.

²¹CHOUDHRI and HAKURA (2006) found that long-run ERPT for Greece and Portugal was equal to 0.42 and 0.44 respectively, after 4 quarter, and 0.48 and 0.54 respectively, after 20 quarter.

6.2. Results from the STR pass-through models

Estimation results for the STR pass-through model are based on equation (4.21). As mentioned above, the parameters of STR model are estimated by NLS estimation technique which provides estimators that are consistent and asymptotically normal. Regarding our choice of transition variable to be included in the final non-linear model, no remaining non-linearity tests are also conducted after estimation. Therefore, we select the transition variables that provided the strongest rejection of both the null of linearity of the baseline linear model, and of no additive non-linearity after estimation of the non-linear model. In choosing the transition function, we employ the sequence of null hypotheses for selecting the STR specification together with the economic intuition. As explained before, the LSTR model is preferred to ESTR model when using inflation rate and output growth as transition variables. And when considering the exchange rate as transition variable, both LSTR and ESTR specification can be used, but we must be careful in our interpretation of the induced dynamic by each specification. The LSTR model captures the pass-through asymmetry during currency appreciations and depreciations episodes, while the ESTR is appropriate to account for non-linearity in pass-through with respect to the size of exchange rate movements. In addition to this, we also gave preference for models that performs well in terms of misspecification tests, i.e. with no error autocorrelation, no additive linearity and with constant parameters.²²

6.2.1. Inflation rate as transition variable

In this section we investigate whether the ERPT responds non-linearly to the inflation level in 12 EA countries. It is argued in the pass-through literature that the responsiveness of prices to exchange rate fluctuations depends positively on inflation. A high inflation environment tends to increase the extent of pass-through. Consequently, we aim to explore this inflation regime-dependence of ERPT in a nonlinear fashion. According to linearity tests (Tables D.7), LSTR model is found to be the best specification to capture this kind of behavior for most of EA countries. This is consistent with theoretical priors that pass-through mechanisms may be different whether inflation rate is above or

²²The highest R^2 and the lowest AIC value are also considered.

below a given threshold. The NLS estimates of our LSTR models are summarized in Table 4.1. We report both short-run and long-run pass-through coefficients as defined in equation (4.22) and (4.23).²³ We compute sum of squared residuals ratio (SSR_{ratio}) between LSTR model and the linear specification which suggests a better fit for the non-linear model. Similarly, the R^2 and the AIC favor the LSTR model against the linear regression. We also check the quality of the estimated LSTR models by conducting several misspecification tests. In most of cases, the selected LSTR models pass the main diagnostic tests, i.e. no error autocorrelation, no conditional heteroscedasticity, parameters constancy and non remaining nonlinearity.

ERPT results in Table 4.1 show significant threshold inflation rate levels for most of the EA countries. Thresholds do not differ considerably across countries. Values are ranging from 1% to 3% with exception of Portugal showing $c = 8\%$. Regarding speed of transition γ , our results indicate relatively moderate values which is a proof of smooth transition between the two inflation regimes.²⁴ When considering short-run ERPT, our results point a significant positive relationship between inflation rates and the extent of pass-through for 5 out of 12 countries. For those 5 EA countries, when inflation increases above the threshold, exchange rate transmission becomes higher. For example, when the Italian CPI inflation exceed 3%, the rate of pass-through increases from 0.03% (when $G = 0$) to about 0.17% (when $G = 1$). For the long-run ERPT, the presence of regime-dependence is more apparent. There are 8 out of 12 EA countries showing a positive link between pass-through and inflation environment. For example, the ERPT in France is equal to 0.08% when inflation rate is below 1%, but beyond this threshold level, ERPT becomes higher and reaches 0.18% (see Figure 4.4).

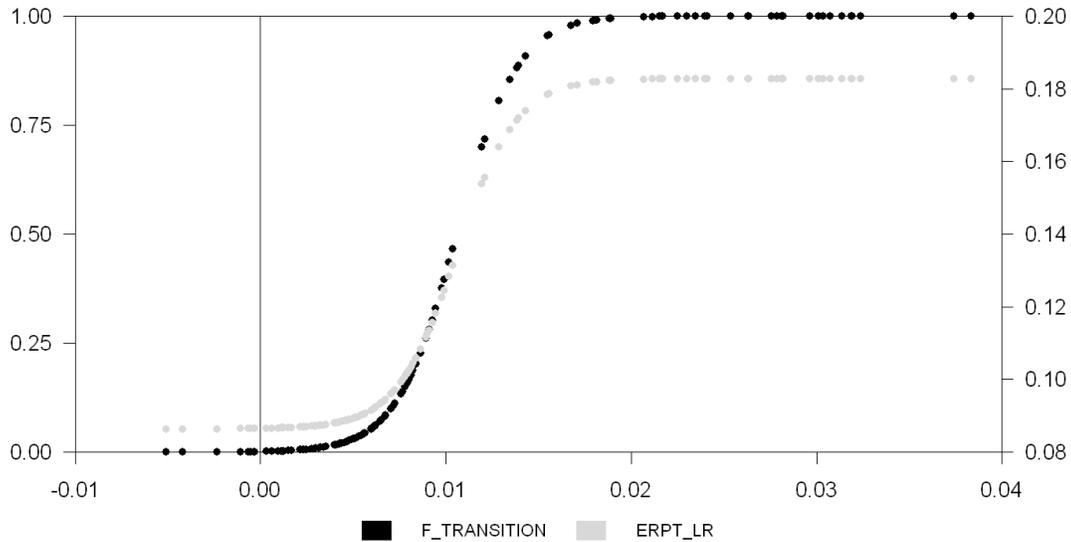
Broadly speaking, our results are in line with Taylor's hypothesis, i.e. the responsiveness of prices to exchange rate fluctuations depends positively on inflation environment. The intuition behind this phenomenon may be due to the foreign firms' behavior. The latter are more willing to set their prices in the currency of importing countries where inflation environment is stable (LCP strategy). In such case ERPT would be lower. But when exporters perceive a higher inflation level, they may shift away from

²³Full results from all STR models are presented in the Tables D.11, D.12, D.13 and D.14 in Appendix D.4.

²⁴According to VAN DIJK, TERÄSVIRTA, and FRANSES (2002) estimates of γ may appear to be insignificant. This should not be interpreted as evidence of weak nonlinearity.

local-currency pricing by passing exchange rate changes through the prices in importer's currency. This behavior would entail a higher degree of pass-through. From empirical point of view, our findings corroborate the scarce ERPT literature using STR models.

Figure 4.4: Estimated transition function and long-run ERPT as a function of past inflation rates in France



As mentioned in section 4, NOGUEIRA JR. and LEON-LEDESMA (2008) has employed LSTR model to capture nonlinearities in pass-through with respect to inflation rate. They conclude that the adoption of inflation target has entailed a lower pass-through for 4 countries in their sample, namely Canada, Mexico, South Africa, and United Kingdom. Similarly, SHINTANI, TERADA-HAGIWARA, and TOMOYOSHI (2013) found that the period of low ERPT is likely to be associated with the low inflation environment in United States, even though the authors used a U-shaped symmetric transition functions instead of an asymmetric logistic function.²⁵

²⁵Two kinds of symmetric transition functions has employed by the authors see section 4 for more details.

Table 4.1: Estimated ERPT elasticities from the LSTR model with $s_t = \pi_{t-i}$

	Austria	Belgium	Germany	Spain	Finland	France	Greece	Ireland	Italy	Luxembourg	Netherlands	Portugal
Transition variable (s_t)	π_{t-4}	π_{t-1}	π_{t-3}	π_{t-4}	π_{t-3}	π_{t-2}	π_{t-3}	π_{t-4}	π_{t-2}	π_{t-4}	π_{t-3}	π_{t-1}
Threshold (c)	0,033 (0,000)	0,030 (0,000)	0,013 (0,000)	0,022 (0,000)	0,027 (0,000)	0,011 (0,000)	0,022 (0,024)	0,034 (0,000)	0,031 (0,000)	0,015 (0,000)	0,008 (0,000)	0,088 (0,000)
Speed of transition (γ)	22,013 (0,547)	17,566 (0,312)	9,390 (0,208)	12,702 (0,437)	13,291 (0,531)	6,134 (0,067)	2,358 (0,120)	8,456 (0,003)	2,449 (0,002)	4,909 (0,056)	9,361 (0,333)	4,061 (0,053)
Linear part : $G = 0$												
<i>SR ERPT</i>	0,043 (0,042)	0,091 (0,000)	0,063 (0,000)	0,085 (0,009)	0,044 (0,005)	0,066 (0,001)	0,105 (0,134)	0,043 (0,097)	0,032 (0,050)	0,053 (0,002)	0,049 (0,009)	0,040 (0,547)
<i>LR ERPT</i>	0,154 (0,112)	0,140 (0,000)	0,115 (0,019)	0,438 (0,246)	0,415 (0,002)	0,086 (0,108)	0,168 (0,478)	0,440 (0,036)	0,183 (0,049)	0,436 (0,112)	0,131 (0,030)	0,059 (0,605)
Non-linear part: $G = 1$												
<i>SR ERPT</i>	0,024 (0,181)	0,075 (0,000)	0,003 (0,969)	0,167 (0,000)	0,020 (0,787)	-0,005 (0,791)	0,056 (0,165)	1,913 (0,001)	0,167 (0,001)	0,159 (0,000)	0,036 (0,066)	0,085 (0,092)
<i>LR ERPT</i>	0,328 (0,027)	0,155 (0,000)	0,251 (0,192)	0,608 (0,205)	0,781 (0,249)	0,183 (0,034)	0,568 (0,103)	2,377 (0,046)	0,904 (0,036)	1,049 (0,138)	0,179 (0,018)	0,492 (0,076)
R^2	0,737	0,757	0,721	0,830	0,818	0,915	0,873	0,803	0,934	0,751	0,727	0,825
SSR_{ratio}	0,786	0,728	0,747	0,798	0,684	0,702	0,842	0,681	0,599	0,778	0,857	0,624
AIC	-8,087	-8,176	-8,338	-6,889	-7,755	-8,536	-6,337	-6,815	-7,940	-8,059	-8,184	-6,052
pJB	0,177	0,146	0,171	0,000	0,003	0,069	0,000	0,000	0,000	0,000	0,967	0,000
$pLM_{AR(4)}$	0,963	0,907	0,083	0,153	0,002	0,136	0,031	0,506	0,616	0,146	0,515	0,248
$pLM_{ARCH(4)}$	0,526	0,204	0,741	0,002	0,747	0,951	0,186	0,439	0,113	0,537	0,586	0,000
pLM_C	0,019	0,028	0,933	0,036	0,164	0,748	0,165	0,000	0,000	0,041	0,183	0,014
pLM_{RNL}	0,361	0,085	0,481	0,027	0,337	0,220	0,590	0,000	0,622	0,578	0,317	0,004

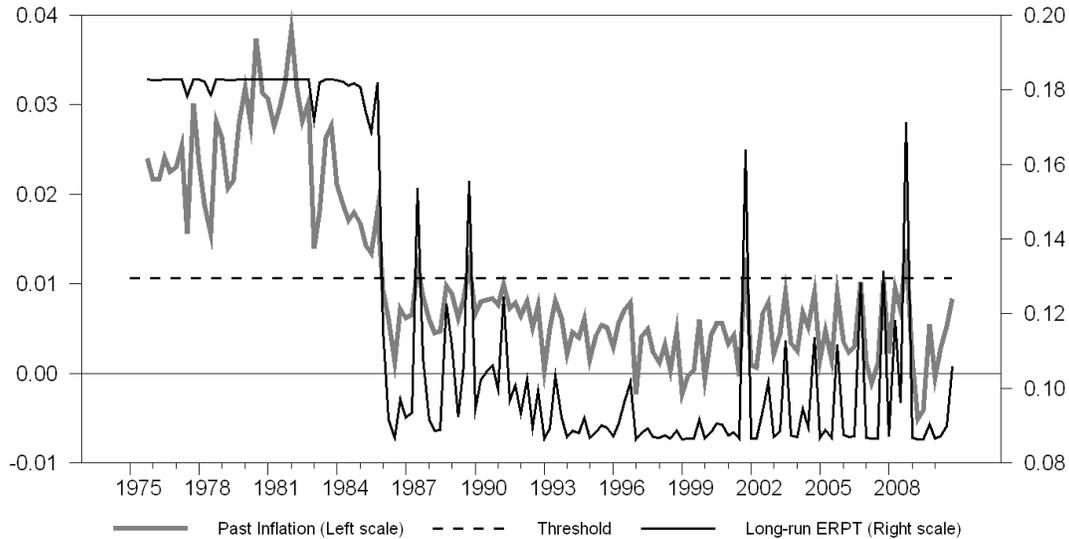
Note: Table reports elasticities of exchange rate pass-through into CPI inflation from LSTR models. Numbers in parentheses are p -values of estimates. R^2 denotes the coefficient of determination, SSR_{ratio} is the ratio of sum of squared residuals between LSTR model and the linear specification, and AIC is the Akaike Information Criterion. The following rows corresponds to the misspecification tests: pJB is the p -values of Jarque-Bera normality test, $pLM_{AR(4)}$ is the p -values of the LM test of no error autocorrelation up to fourth order, $pLM_{ARCH(4)}$ is the p -values of the LM test of no ARCH effects up to fourth order, pLM_C is the p -values of the LM test of parameter constancy and pLM_{RNL} is the p -values of the LM test of no remaining nonlinearity.

Additionally, we have plotted both the estimated transition functions and the ERPT as a function of the transition variable lagged inflation ($s_t = \pi_{t-i}$). Graphs of both short- and long-run pass-through are presented respectively in Figure D.1 and D.2 in Appendix D.5.²⁶ It is clear that the transition between both extreme regimes, i.e. $G(s_t; \gamma, c) = 0$ and $G(s_t; \gamma, c) = 1$, is smooth in most of cases. Plots reveal the regime dependence of ERPT to inflation environment. The positive connection between the degree of the ERPT and inflation is quite clear, except for Belgium and Netherlands where the relationship is negative in the short-run.

To give further insight of inflation regime dependence, we plot the time-variation of ERPT over inflationary and disinflationary episodes between 1975-2010. Results of time-varying pass-through coefficients are given in Figure D.11 and D.12 in Appendix D.6. We also report lagged inflation rates and the estimated threshold level of inflation on the same graph. A careful inspection of the plots shows that the exchange rate transmission was higher during the second half of the 1970s and the early of 1980s for most of EA countries. Over this period, there had been an unstable inflation environment due to the oil shocks of the 1970s. As shown in Figure D.11 and D.12, inflation rates was exceeding the threshold levels in our country sample which has resulted in increased degree of pass-through during this episode.

It is worth noting that since the late 1980s and the beginning of 1990s, most of EA countries has entered an era of low inflation regime (see Figure 4.5 for the French case). According to Figure D.11 and D.12, this shifting towards stable inflation has coincided with the decline of the extent of pass-through. The bulk of recent literature of pass-through has documented this lowering of the domestic price sensibility to exchange rate variation, including BAILLIU and FUJII (2004), GAGNON and IHRIG (2004). Another important remark is that the low-inflation regime is more recent for Greece and Portugal compared to the rest of our country sample, i.e. late of 1990s for Greece and mid of the 1990s for Portugal. This may explain why pass-through estimates based on linear model (equation (4.28)) are higher in Greece and Portugal in comparison with the rest of EA countries. During our sample period (1975-2010), there was unstable inflation environment for these two countries and this helps explain their relatively large rate of ERPT.

²⁶We only report results for countries with significant coefficient of pass-through.

Figure 4.5: Time-varying long-run ERPT and past inflation in France

Note: Time-varying long-run ERPT and past inflation during 1975-2010. Results are from LSTR model with $s_t = \pi_{t-i}$.

6.2.2. Exchange rate as transition variable

In this section, we consider the rate of exchange rate depreciation (Δe_{t-i}) as the driving factor of the nonlinearity. As mentioned above, there are two types of linearities that can be modeled. On one hand, pass-through asymmetries can arise with respect to exchange rate change direction i.e., in response to currency depreciations and appreciations. We have seen that LSTR specification is pertinent in situations in which the dynamic behavior is different whether the transition variable is below or above the threshold. Therefore, we employ LSTR model to capture ERPT asymmetry during currency appreciations and depreciations episodes, especially when the threshold level of Δe_{t-i} is close to zero. On the other hand, there is a second type of linearities which is related to the size of exchange rate movements. If firms are willing to absorb small changes in exchange rate rather than larger ones due to the presence of "menu costs", this may result in asymmetric pass-through of large and small exchange rate shocks. In such case, ESTR specification would be more appropriate in describing this non-linearity in ERPT mechanism.

As regards linearity tests as reported in Table D.8, there is an evidence of presence of nonlinearity in all EA countries except for Austria. Once linearity has been rejected, we employ the sequence of null hypotheses for selecting the relevant transition function, i.e. logistic or exponential. As discussed in VAN DIJK, TERÄSVIRTA, and FRANSES (2002), recent increases in computational power have made the decision rule, based on testing a sequence of nested null hypotheses, less important in practice. The authors argued that is easy to estimate a number of both LSTAR and ESTAR models and choose between them at the evaluation stage by misspecification tests. In addition, economic intuition must be considered in selecting the adequate transition function. In their study, NOGUEIRA JR. and LEON-LEDESMA (2008) examined the role of the size of the exchange rate movements in generating asymmetry by implementing an ESTR specification. However, in our work, we aim to explore the two possible sources of nonlinearities in ERPT, i.e. with respect to both direction and magnitude of exchange rate changes. Therefore, we follow VAN DIJK, TERÄSVIRTA, and FRANSES (2002) approach by estimating a number of both LSTAR and ESTAR models for each country. This is a sensible way to check what kind of asymmetry that really drives the nonlinear mechanism in ERPT. As explained in section 3.1, LSTR specification would be more appropriate to capture asymmetry arising from the *direction* of exchange rate changes, while ESTR specification is more suitable for asymmetric behavior with respect to the size of exchange rate movements.

Results from LSTR model As summarized in Table 4.2, we report only results for countries with significant ERPT coefficient. As we can see, the nonlinear models provides a better fit to the data than the linear models with respect to R^2 , SSR and AIC . We note that there are only 5 out of 12 EA countries show a significant response of CPI inflation to exchange rate movements in a nonlinear way. The threshold levels are very close for Italy, Luxembourg and Portugal (around 4%), but differ greatly in comparison to Belgium and Greece. The same thing for the speed of transition which varies across those countries.²⁷ Concerning ERPT estimates, our results are to some extent mixed. For Italy, Luxembourg and Portugal, when exchange rate is depreciating above some threshold level, the short-run pass-through becomes higher. For example,

²⁷We note that the parameters γ is very high in Belgium, which indicates an abrupt transition rather than a smooth one.

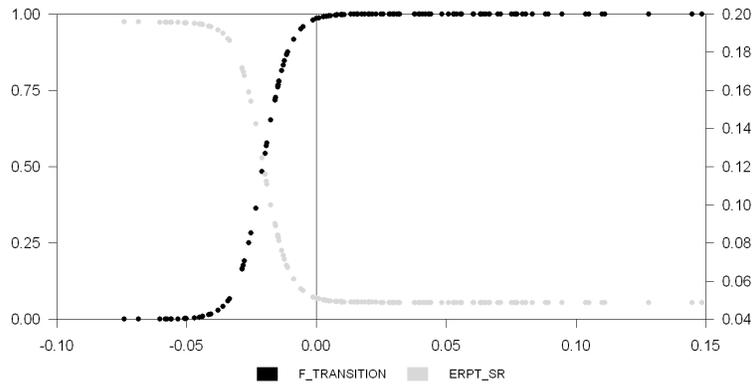
short-run ERPT coefficient rise from 0.07% to 0.27% in Portugal once the rate of currency depreciation is exceeding 4.5% (see Figure 4.6). We can say that exchange rate transmission is higher for large rate of depreciation, but it becomes lower for small depreciation and in appreciation episodes.²⁸ These results seem to be consistent with the so-called *capacity constraints* hypothesis. Since quantities may be rigid upwards in the short run, exporters may not be able to increase sales when importing country currency is appreciating. So, they are willing to raise markup and let quantity unchanged. In this case, pass-through would be greater when the importer's currency is appreciating than when it is depreciating.

Table 4.2: Estimated ERPT elasticities from the LSTR model with $s_t = \Delta e_{t-i}$

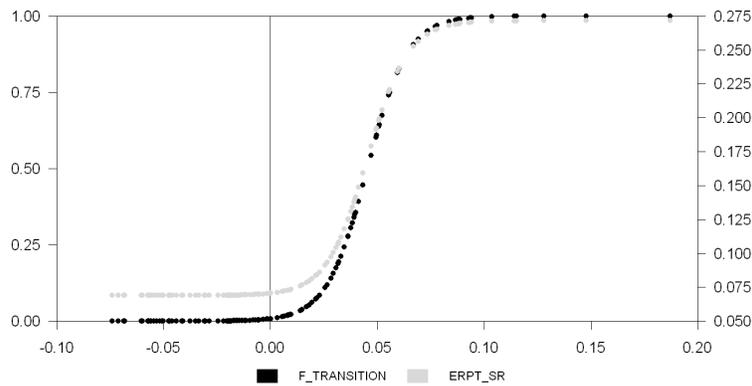
	Belgium	Greece	Italy	Luxembourg	Portugal
Transition variable (s_t)	Δe_{t-4}	Δe_{t-4}	Δe_{t-2}	Δe_{t-1}	Δe_{t-1}
Threshold (c)	0,004 (0,050)	-0,021 (0,000)	0,044 (0,000)	0,037 (0,000)	0,045 (0,000)
Speed of transition (γ)	60,750 (0,555)	9,675 (0,262)	7,513 (0,095)	18,530 (0,379)	5,317 (0,029)
Linear Part : $G = 0$					
<i>SR ERPT</i>	0,101 (0,000)	0,196 (0,033)	0,036 (0,030)	0,060 (0,000)	0,069 (0,131)
<i>LR ERPT</i>	0,285 (0,000)	0,518 (0,256)	0,433 (0,000)	0,176 (0,000)	0,101 (0,564)
Non-linear part : $G = 1$					
<i>SR ERPT</i>	0,041 (0,016)	0,049 (0,081)	0,101 (0,106)	0,123 (0,001)	0,272 (0,000)
<i>LR ERPT</i>	0,151 (0,006)	0,442 (0,299)	-0,107 (0,780)	0,201 (0,052)	2,029 (0,000)
R^2	0,723	0,904	0,911	0,751	0,805
SSR_{ratio}	0,828	0,634	0,803	0,778	0,694
AIC	-8,047	-6,531	-7,648	-8,059	-5,976
pJB	0,718	0,000	0,000	0,026	0,001
$pLM_{AR(4)}$	0,436	0,094	0,977	0,876	0,315
$pLM_{ARCH(4)}$	0,625	0,440	0,008	0,867	0,005
pLM_C	0,165	0,303	0,020	0,137	0,012
pLM_{NLR}	0,069	0,154	0,548	0,416	0,168

Note: Table reports elasticities of exchange rate pass-through into CPI inflation from LSTR models. Numbers in parentheses are p -values of estimates. R^2 denotes the coefficient of determination, SSR_{ratio} is the ratio of sum of squared residuals between LSTR model and the linear specification, and AIC is the Akaike Information Criterion. The following rows corresponds to the misspecification tests: pJB is the p -values of Jarque-Bera normality test, $pLM_{AR(4)}$ is the p -values of the LM test of no error autocorrelation up to fourth order, $pLM_{ARCH(4)}$ is the p -values of the LM test of no ARCH effects up to fourth order, pLM_C is the p -values of the LM test of parameter constancy and pLM_{NLR} is the p -values of the LM test of no remaining nonlinearity.

²⁸We have the same pattern in the long run for Luxembourg and Portugal (See Figure D.4 in Appendix D.5).

Figure 4.6: Estimated logistic function and short-run ERPT as a function of past depreciations

Greece



Portugal

On the other hand, when we look to the ERPT estimates for Belgium and Greece, results are quietly different. The short-run response of CPI inflation to exchange rate is negatively correlated with the rate of depreciation (See Figure 4.6 for the Greek case).²⁹ For Belgium, threshold level is close to zero ($c = 0.004$), and we can say that short-run ERPT decreases significantly from 0.1% to 0.04% as the exchange rate is depreciating. As a result, the extent of pass-through is smaller during the depreciation

²⁹The same thing is found for Belgium in the long run (see Figure D.4).

than in appreciation episodes. This is in line with the thesis of *Market share objective*. Faced with a depreciation of the importing country's currency, foreign firms can follow pricing-to-market strategy by adjusting their markups to maintain market. But in the case of an appreciation, they maintain their markups and allow the import price to fall in the currency of destination market. Consequently, an appreciation of the importing country's currency might cause larger pass-through than depreciation.

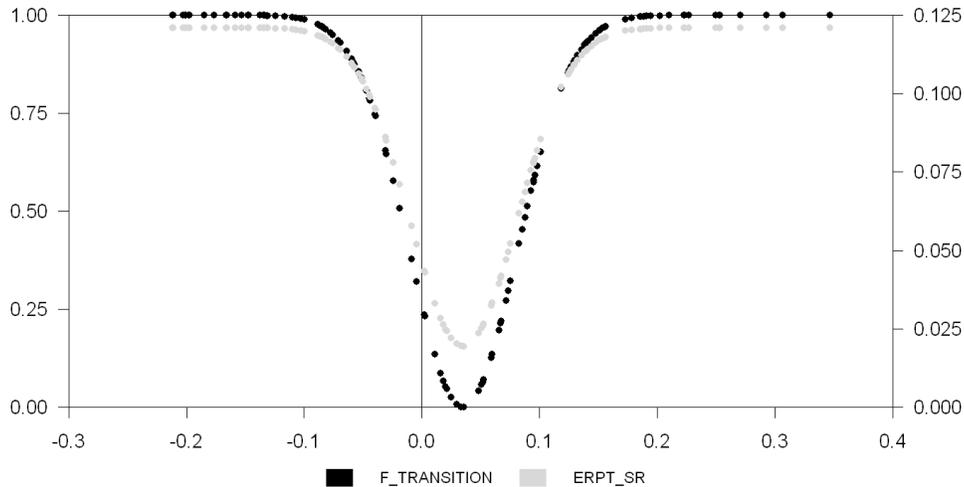
In all, our results are somewhat mixed since there is no clear direction of asymmetry. For 3 out of 5 EA countries (Italy, Luxembourg and Portugal), ERPT is greater when exchange rate is depreciating, while for Belgium and Greece, pass-through is lower when importer's currency is depreciating. Nevertheless, our findings corroborate with previous empirical studies which provide also no clear evidence on the direction of asymmetry in ERPT. For a set of European industries, GIL-PAREJA (2000) found that the direction of the asymmetry varied across industries and countries. COUGHLIN and POLLARD (2004) confirm the same results in their study on 30 U.S. industries.

Results from ESTR model The second type of possible linearity is related to the magnitude of exchange rate change. The extent of pass-through may respond asymmetrically to the size of currency fluctuations, in the sense that there is differential effect of large versus small exchange rate shocks. As discussed above, an ESTR specification would be more appropriate to capture this kind of asymmetric behavior. Although ESTR model allow for symmetric dynamics with respect to negative and positive deviations of exchange rate changes from the threshold level, the mechanism would be asymmetric depending on whether Δe_{t-i} is close or far away from the threshold c . In other words, what matters here is the size exchange rate movements.

In Table 4.3, we report only countries with significant pass-through elasticity. As we can see, most of EA countries (except Austria and Portugal) exhibit a significant nonlinear response of CPI inflation to exchange rate movements. Especially in the short-run, there is 9 EA countries with an evidence of positive correlation between pass-through and the magnitude of currency changes. In Spain, the short-run ERPT coefficient is not statistically significantly different from zero when exchange rate variation is small

- when Δe_{t-i} is close to the threshold of $c = 0.006$. But for large currency movements, i.e. when Δe_{t-i} is far away from the threshold level, the Spanish short-run pass-through corresponds to 0.12% (see Figure 4.7).

Figure 4.7: Exponential functions and short-run ERPT as a function of past depreciations in Spain



Figures D.5 and D.6 in Appendix D.5 give a supportive evidence of the presence of asymmetries arising from the size of exchange rate shocks. That is, large exchange rate changes elicit greater ERPT. This result is consistent with the *menu costs* assumption. If foreign firms perceive that price changes are costly, a small currency change can be accommodated within the mark-up. But, if exchange rate changes exceed some threshold, firms are tempted to change their prices in the currency of importing country. Empirically, NOGUEIRA JR. and LEON-LEDESMA (2008) has put forth the role of menu costs in explaining nonlinearities in ERPT. To the best of our knowledge, it is the only work using ESTR model in this context. The results of NOGUEIRA JR. and LEON-LEDESMA (2008) suggest that only two (Mexico and UK) out of six countries provide an evidence of non-linear ERPT with respect to the size of exchange rate changes.

Table 4.3: Estimated ERPT elasticities from the ESTR model with $s_t = \Delta e_{t-i}$

	Belgium	Germany	Spain	Finland	France	Greece	Ireland	Italy	Luxembourg	Netherlands
Transition variable (s_t)	Δe_{t-4}	Δe_{t-1}	Δe_{t-4}	Δe_{t-2}	Δe_{t-3}	Δe_{t-3}	Δe_{t-2}	Δe_{t-1}	Δe_{t-3}	Δe_{t-4}
Threshold (c)	0,022 (0,059)	0,006 (0,037)	0,035 (0,004)	0,021 (0,000)	-0,022 (0,000)	0,030 (0,000)	0,043 (0,000)	0,016 (0,000)	0,010 (0,016)	0,033 (0,000)
Speed of transition (γ)	4,381 (0,000)	11,092 (0,062)	4,322 (0,110)	11,347 (0,004)	2,487 (0,064)	33,264 (0,053)	1,274 (0,025)	9,112 (0,105)	4,041 (0,057)	1,128 (0,058)
Linear part : $G = 0$										
<i>SR ERPT</i>	-0,016 (0,681)	0,002 (0,972)	0,019 (0,814)	-0,071 (0,183)	-0,019 (0,485)	-0,291 (0,073)	0,065 (0,256)	0,009 (0,886)	0,055 (0,062)	-0,018 (0,476)
<i>LR ERPT</i>	-0,107 (0,648)	-0,478 (0,516)	-0,286 (0,582)	-0,129 (0,327)	0,036 (0,807)	0,414 (0,441)	0,125 (0,164)	-0,211 (0,844)	0,090 (0,228)	0,030 (0,908)
Non-linear part: $G = 1$										
<i>SR ERPT</i>	0,103 (0,000)	0,075 (0,000)	0,121 (0,000)	0,050 (0,004)	0,077 (0,000)	0,104 (0,000)	-0,010 (0,750)	0,070 (0,000)	0,090 (0,000)	0,104 (0,000)
<i>LR ERPT</i>	0,573 (0,044)	0,435 (0,488)	0,671 (0,176)	0,122 (0,003)	0,374 (0,042)	-0,237 (0,378)	0,215 (0,008)	0,465 (0,239)	0,265 (0,007)	1,192 (0,572)
R^2	0,660	0,573	0,787	0,796	0,884	0,882	0,802	0,902	0,742	0,737
<i>SSR</i>	0,826	1,147	1,020	0,768	0,962	0,781	0,685	0,886	0,807	0,827
<i>AIC</i>	-8,050	-7,969	-6,674	-7,639	-8,167	-6,412	-6,779	-7,519	-8,023	-8,221
<i>pJB</i>	0,229	0,000	0,000	0,035	0,002	0,450	0,000	0,000	0,132	0,464
$pLM_{AR(4)}$	0,454	0,000	0,582	0,043	0,000	0,000	0,123	0,147	0,834	0,850
$pLM_{ARCH(4)}$	0,340	0,801	0,521	0,010	0,640	0,000	0,154	0,389	0,224	0,293
pLM_C	0,070	0,605	0,137	0,131	0,166	0,450	0,456	0,037	0,253	0,207
pLM_{NLR}	0,113	0,199	0,370	0,368	0,572	0,659	0,107	0,328	0,220	0,253

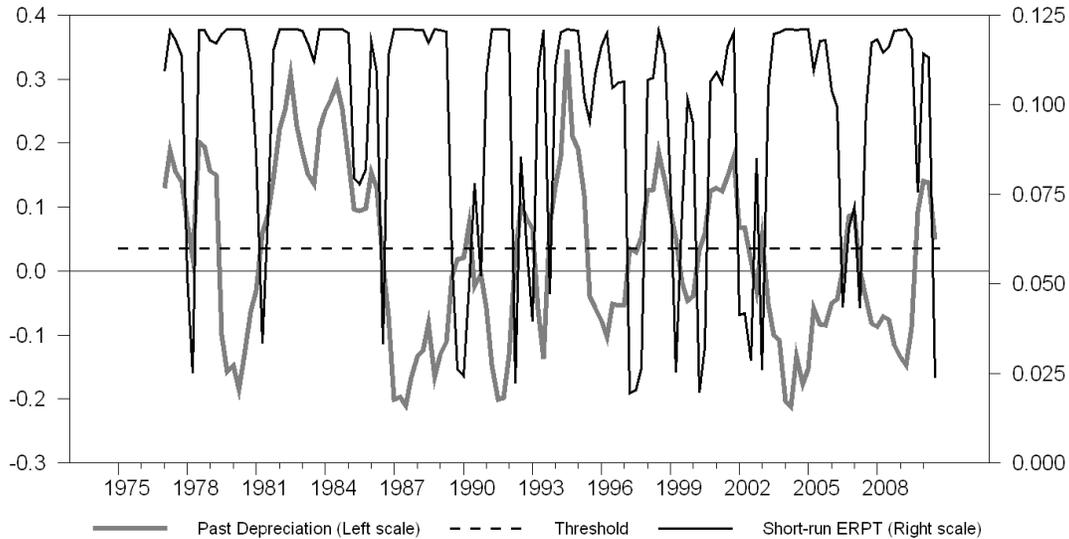
Note: Table reports elasticities of exchange rate pass-through into CPI inflation from LSTR models. Numbers in parentheses are p -values of estimates. R^2 denotes the coefficient of determination, SSR_{ratio} is the ratio of sum of squared residuals between ESTR model and the linear specification, and AIC is the Akaike Information Criterion. The following rows corresponds to the misspecification tests: pJB is the p -values of Jarque-Bera normality test, $pLM_{AR(4)}$ is the p -values of the LM test of no error autocorrelation up to fourth order, $pLM_{ARCH(4)}$ is the p -values of the LM test of no ARCH effects up to fourth order, pLM_C is the p -values of the LM test of parameter constancy and pLM_{NLR} is the p -values of the LM test of no remaining nonlinearity.

Concerning the evolution of ERPT over time, plots are reported in Figure D.15 and D.16. An interesting result concerns the period of launching the euro area. It is well-known that EA countries - except Greece that joined the monetary union in 2001 - have experienced an ongoing depreciation of the euro between the end of 1998 until the last quarter of 2001.³⁰ While since the mid-2002 the euro has started a steady appreciation until the end of 2004. As argued by BUSSIÈRE (2012), such dramatic changes in the value of European currency may give rise to asymmetric pass-through. Thereby, it is clear from the visualization of Figure D.15 that ERPT was higher following the introduction of the euro for most of our EA countries (see Figure 4.8 for the Spanish case). According to our results, the dramatic change of the European currency during the first three years of the euro has elicited a higher rate of pass-through. When the depreciation of the euro surpassed some limit, those countries have experienced a higher response of CPI inflation which can be interpreted as a proof of nonlinear mechanisms of pass-through.

Also, another prominent result is relative to the European Monetary System (EMS) crisis (1992-1993). During this episode, the extent of pass-through was higher for most of EA countries. It is known that for members of EMS, currencies were allowed to fluctuate within pre-specified bands (a system known as the Exchange Rate Mechanism (ERM)). During the crisis period, a wave of devaluations has occurred for major EMS countries, especially for Italy that was forced to withdraw the ERM in September 1992.³¹ Consequently, due to the excessive variability of the European currencies (conjugated with confidence crisis), it is expected that foreign firms tend to modify pricing strategy, shifting from importer's currency pricing (*LCP* strategy) to exporter's currency invoicing (*PCP* strategy). As a result, the degree of pass-through is found to be higher during this episode. Similarly, one might say that the EMS crisis could be an illustration of asymmetric mechanisms of ERPT with respect to the magnitude of exchange rate change. When exchange rate changes surpass some limit, the exchange rate transmission becomes larger.

³⁰During this period, the euro has depreciated by nearly 20% in nominal effective terms.

³¹Austria, Finland and Greece were not member of the ERM at that time.

Figure 4.8: Time-varying short-run ERPT and past depreciations in Spain

Note: Time-varying short-run ERPT and past depreciations during 1975-2010. Results are from ESTR model with $s_t = \Delta e_{t-j}$.

6.2.3. Output growth as transition variable

In this part of our analysis, we raise the question of whether the degree of ERPT is affected by the business cycle in a nonlinear way. The sparse empirical evidence on this issue has put forth a positive relationship between economic activity and the transmission of exchange rate. Intuitively, in periods where the economy is booming, firms are more willing to pass-through cost increases such as those coming from the exchange rate, meaning that ERPT would be greater in periods of prosperity than in periods of slowdown. In accordance with this argument, GARCÍA and RESTREPO (2001) has explained that the lower ERPT in Chile in the 1990s is due, in part, to the positive dependence of pass-through to economic activity. According to the authors, the negative output gap during this period has offset the inflationary impact of exchange rate depreciation by reducing margins. Furthermore, as is well-known, markups and profit margins are pro-cyclical. Then, prices would move in the same direction with the business cycle, increasing during the expansion and decreasing during economic slowdown. Also, the power of wage negotiations is more important during recovery

periods, which may lead to price increases. Thereby, exporters are more willing to pass currency changes through prices when the economy is booming than in periods of slowdown.

The asymmetric reaction of ERPT over the business cycle was confirmed by GOLDFAJN and WERLANG (2000) in a panel of 71 countries. The authors have found that depreciations have a higher pass-through to prices during prosperity periods. In a Phillips curve threshold framework, CORREA and MINELLA (2006) and PRZYSTUPA and WRÓBEL (2011) suggest that when the output gap is above a certain threshold, ERPT becomes higher. To the best of our knowledge, only the study of NOGUEIRA JR. and LEON-LEDESMA (2008) that used LSTR model to capture nonlinearity in ERPT with respect to the business cycle. The authors investigated the presence of nonlinearities in a sample of 6 developed and developing Inflation Target countries.

In our analysis, we follow NOGUEIRA JR. and LEON-LEDESMA (2008) approach by using LSTR specification which can describe an asymmetric behavior depending on whether the transition variable is below or above the threshold. The economic activity is considered as the driving factor of the nonlinear dynamic. As a proxy for the economic activity along the business cycle, we consider the rate of growth of the real GDP.³² Thus, the lagged real GDP growth is considered as the transition variable ($s_t = \Delta y_{t-j}$) in the STR model. When its values exceeding an estimated threshold, these can be interpreted as periods of expansion. While, when values are below the threshold, these are periods of economic slowdown or recession. The choice of the adequate lagged real GDP growth as a transition variable by means of linearity tests is reported in Table D.9 in Appendix D.3. According to linearity tests, there is a strong evidence of presence of nonlinearities in 9 out of 12 EA countries (except for France, Ireland and Luxembourg). As explained before, the economic intuition must be also considered in our choice of the relevant STR specification. According to VAN DIJK, TERÄSVIRTA, and FRANSES (2002), LSTR models are more appropriate in describing processes whose dynamic properties are different in expansions from what they are in recessions. Effectively, in accordance with theoretical priors (section 2.2), the ERPT may be different whether

³²In their studies, GOLDFAJN and WERLANG (2000) and CORREA and MINELLA (2006) used the output gap as proxy for the economic activity. However, as explained by NOGUEIRA JR. and LEON-LEDESMA (2008), the use of an ad hoc detrending processes like the output gap might eliminate valuable information from the data.

economic activity is above or below a given threshold. In other words, the exchange rate changes would have a higher pass-through when the economy is growing faster than when the output growth is below the threshold. Thus, given these features, the LSTR model is preferred to ESTR. Estimation results from the LSTR pass-through equation (4.21) are summarized in Table 4.4. They concern only EA countries rejecting the null of linearity (9 out of 12 EA countries). In addition to the estimated threshold level and the speed of transition, we report ERPT coefficients for the two extremes regimes, i.e. $G(s_t; \gamma, c) = 0$ and $G(s_t; \gamma, c) = 1$ (low and high activity regimes respectively).³³

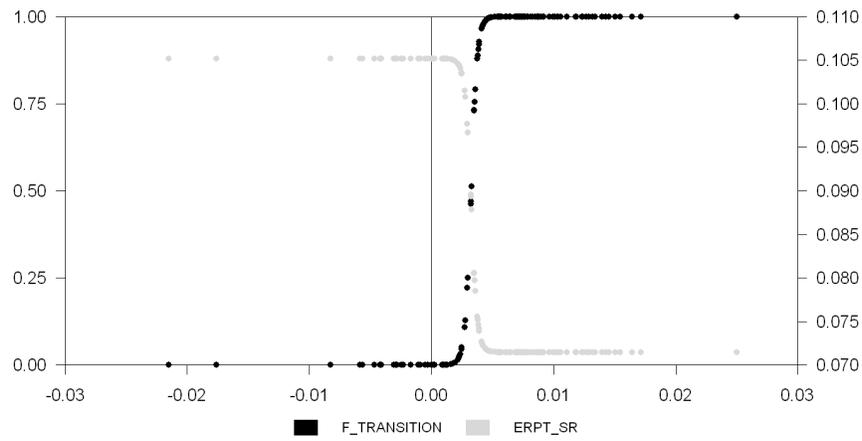
Estimated short- and long-run ERPT from LSTR model are summarized in Table 4.4. From statistical point of view, the model performs well in terms of the goodness of fit and according to misspecification tests. We see that the threshold level of real GDP growth varies significantly across countries, ranging from 0.3% in Belgium to 4% in Austria. Regarding pass-through estimates, there are 6 out of 9 EA countries showing significant nonlinear ERPT with respect to business cycle. In other words, pass-through elasticity is significantly different between low and high activity regimes in 6 EA countries. For these countries, we denote that the extent of pass-through depends positively on economic activity, except for Belgium and Netherlands. For these countries, the exchange rate transmission to CPI inflation is significantly greater when output growth is above some threshold. For instance, the pass-through coefficient in Germany is 0.02% not significantly different from zero when GDP growth is below 1%, i.e. during economic slowdown. However, when German economy is growing faster, above the threshold of 1%, ERPT elasticity increase to about 0.13% (see Figure 4.9).

Also, we have plotted both the estimated transition functions and the ERPT as a function of the transition variable lagged real GDP (see Figures D.7 in Appendix D.5). Plots reveal the regime-dependence of ERPT to business cycle. The positive connection between the degree of the ERPT and real GDP growth is quite clear for 4 out of 6 EA countries in the short-run. These results are consistent with the existing empirical literature dealing with the issue of nonlinearity. In their LSTR model, NOGUEIRA JR. and LEON-LEDESMA (2008) found the same positive link between pass-through and economic activity. This is true for 3 out of their 6 Inflation Target countries. Similarly, in a Phillips curve threshold framework, CORREA and MINELLA (2006) suggest that when

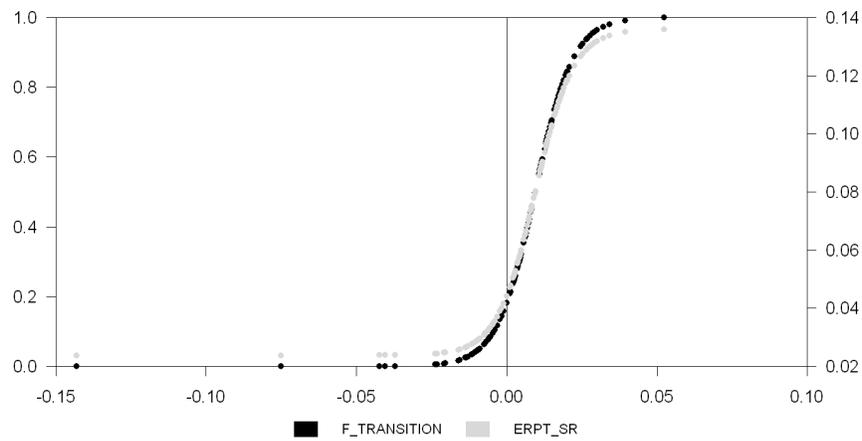
³³Full results of NLS estimates of our LSTR models are presented in the Table D.14 in Appendix D.4.

the output gap is above a certain threshold, ERPT becomes higher in Brazil. Moreover, GOLDFAJN and WERLANG (2000) provide an evidence of asymmetric behavior of ERPT over the business cycle in a panel of 71 countries. The authors found that depreciations have a higher pass-through to prices during prosperity periods.

Figure 4.9: Logistic function and short-run ERPT as a function of past output growth



Belgium



Germany

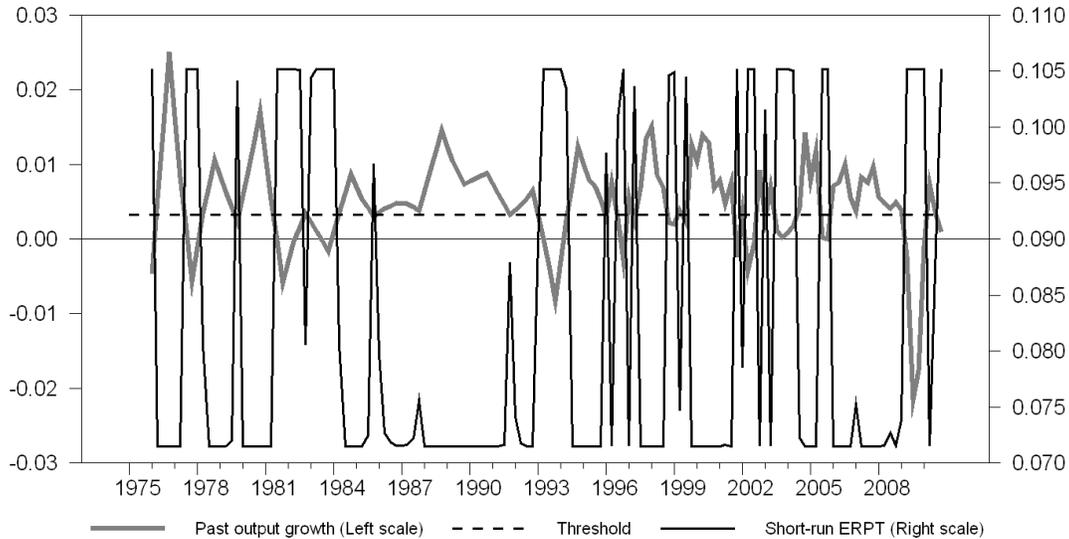
Table 4.4: Estimated ERPT elasticities from the LSTR model with $s_t = \Delta y_{t-i}$

	Austria	Belgium	Germany	Spain	Finland	Greece	Italy	Netherlands	Portugal
Transition variable (s_t)	Δy_{t-1}	Δy_{t-3}	Δy_{t-4}	Δy_{t-3}	Δy_{t-2}	Δy_{t-2}	Δy_{t-1}	Δy_{t-4}	Δy_{t-3}
Threshold (c)	0,040 (0,000)	0,003 (0,000)	0,010 (0,079)	0,006 (0,509)	0,029 (0,000)	0,021 (0,009)	0,017 (0,000)	0,007 (0,000)	0,013 (0,000)
Speed of transition (γ)	24,444 (0,651)	20,760 (0,168)	3,304 (0,162)	26,210 (0,000)	3,740 (0,193)	4,585 (0,202)	3,944 (0,003)	8,959 (0,265)	26,378 (0,311)
Linear part : $G = 0$									
<i>SR ERPT</i>	0,044 (0,001)	0,105 (0,000)	0,024 (0,269)	0,049 (0,129)	0,010 (0,708)	0,112 (0,001)	0,044 (0,000)	0,043 (0,025)	0,093 (0,021)
<i>LR ERPT</i>	0,191 (0,015)	0,328 (0,000)	0,088 (0,135)	0,198 (0,400)	0,148 (0,121)	0,581 (0,001)	0,328 (0,000)	0,208 (0,000)	0,707 (0,000)
Non-linear part: $G = 1$									
<i>SR ERPT</i>	0,222 (0,012)	0,071 (0,000)	0,136 (0,005)	0,163 (0,000)	0,080 (0,007)	0,006 (0,936)	0,073 (0,736)	0,032 (0,075)	0,126 (0,162)
<i>LR ERPT</i>	0,337 (0,250)	0,197 (0,000)	0,180 (0,163)	1,061 (0,116)	0,471 (0,009)	0,279 (0,443)	-1,358 (0,332)	0,156 (0,014)	1,619 (0,042)
R^2	0,735	0,772	0,695	0,845	0,790	0,870	0,954	0,737	0,793
SSR_{ratio}	0,812	0,681	0,818	0,729	0,790	0,859	0,413	0,826	0,736
AIC	-8,087	-8,158	-8,247	-6,979	-7,610	-6,317	-8,311	-8,221	-5,857
pJB	0,466	0,364	0,081	0,000	0,108	0,005	0,000	0,462	0,000
$pLM_{AR(4)}$	0,1898	0,968	0,429	0,393	0,015	0,057	0,543	0,691	0,121
$pLM_{ARCH(4)}$	0,446	0,996	0,058	0,093	0,228	0,316	0,000	0,917	0,019
pLM_C	0,193	0,176	0,625	0,010	0,642	0,088	0,539	0,660	0,241
pLM_{RNL}	0,410	0,851	0,943	0,618	0,787	0,164	0,572	0,506	0,730

Note: Table reports elasticities of exchange rate pass-through into CPI inflation from LSTR models. Numbers in parentheses are p -values of estimates. R^2 denotes the coefficient of determination, SSR_{ratio} is the ratio of sum of squared residuals between LSTR model and the linear specification, and AIC is the Akaike Information Criterion. The following rows corresponds to the misspecification tests: pJB is the p -values of Jarque-Bera normality test, $pLM_{AR(4)}$ is the p -values of the LM test of no error autocorrelation up to fourth order, $pLM_{ARCH(4)}$ is the p -values of the LM test of no ARCH effects up to fourth order, pLM_C is the p -values of the LM test of parameter constancy and pLM_{RNL} is the p -values of the LM test of no remaining nonlinearity.

Nevertheless, in the long run, the positive relationship between ERPT and business is present in only three EA countries (Spain, Finland and Portugal), while for Belgium and Netherlands we have a negative connection as in the short-run. For the latter countries, when real GDP growth is below some threshold, the extent of ERPT becomes higher (see Figure 4.9 for the Belgian case). In fact, this is not surprising if low or negative output growth is seen as a period of economic slump or macroeconomic instability. If foreign producers expect less stable conditions in importing country, they may shift away from local-currency pricing strategy (*LCP* strategy), leaving their prices affected by exchange rate changes. As a result, ERPT would be higher in periods of macroeconomic distress than in prosperity episodes. To give further insight on this plausible negative relationship, we plot time-varying ERPT coefficients over the period 1975-2010 (see Figure D.17 in Appendix D.6). According to Figures, we note that extent of pass-through was higher in both Belgium and Netherlands during periods of contraction or recession. For example, we find an increasing rate of ERPT over the European Monetary System (EMS) crisis (1992-1993) and in the 2008 financial crisis. Due to macroeconomic instability episodes, it is more likely that foreign firms tend to modify pricing strategy by choosing the exporter's currency invoicing (*PCP* strategy) in stead of the importer's currency pricing (*LCP* strategy). Therefore, it is not really surprising that pass-through would be greater in Belgium and Netherlands during these periods.

In all, our results reveal no clear direction in this regime-dependence of ERPT to business cycle. In some countries, ERPT is higher during periods of expansion than in periods of recession; however, in other countries, this result is reversed. The responsiveness of CPI inflation to exchange rate changes along the business cycle is different between these two groups of EA countries. So, we can conclude that the nonlinear mechanism of ERPT with respect to the economic activity is an heterogeneous phenomenon across monetary union members. This outcome would have important implications for the design of monetary policy and the expectation of inflation in the euro area. Monetary policy during turbulent exchange rate periods should factor in the nonlinear mechanism of ERPT over the business cycle and how it affects inflation dynamics.

Figure 4.10: Time-varying short-run ERPT and past output growth in Belgium

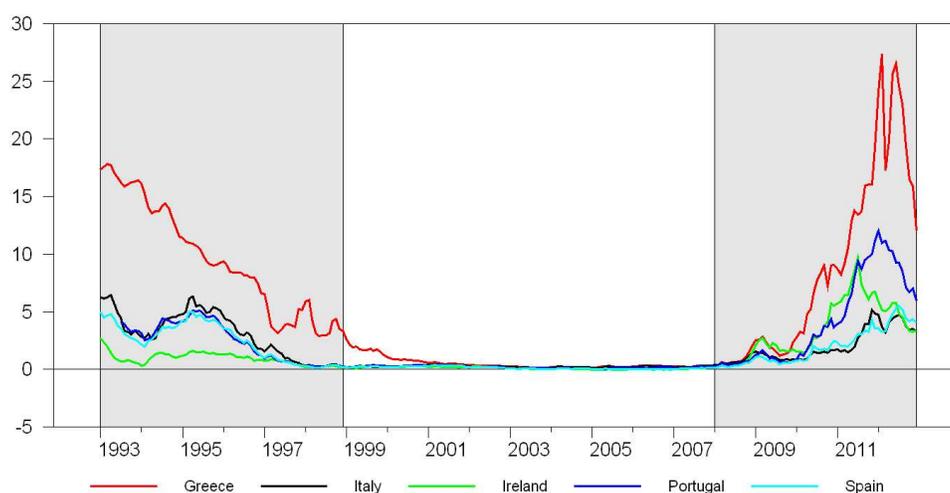
Note: Time-varying short-run ERPT and past depreciations during 1975-2010. Results are from LSTR model with $s_t = \Delta y_{t-i}$.

6.2.4. Sovereign bond yield spread as transition variable

Due to the mixed results concerning the nonlinearity of ERPT with respect to economic activity, we propose an alternative indicator that reflects more accurately the macroeconomic environment of an EA country. As argued by NOGUEIRA JR. and LEON-LEDESMA (2011), exporter's markup depends on the importing country's general macroeconomic stability. The firm's decision on how much to pass-through exchange rate movements into prices depends on its view on the importing country's macroeconomic conditions. When the economy faces a financial or a confidence crisis foreign firms may decide to pass-through a larger proportion of its cost changes in view of the increased likelihood of default from the importer. However, in periods of good macroeconomic conditions, prices will become more insulated from exchange rate changes since foreign firms are willing to adopt local pricing strategy (LCP). This intuition is in line with Taylor's hypothesis, i.e. countries with low stable monetary policies are more likely to have their currencies chosen for transaction invoicing, and hence more likely to have low pass-through to domestic prices.

Therefore, in our empirical analysis we must look for a suitable proxy for macroeconomic stability/instability. In their LSTR model, NOGUEIRA JR. and LEON-LEDESMA (2011) used the real interest rate differential of Mexico with respect to the U.S. as measures of macroeconomic instability, which is the transition variable in the nonlinear smooth transition model.³⁴ The use of real interest rate spread as a proxy of macroeconomic instability, and particularly as a leading indicator of confidence crises, has been advocated, among others, by KAMINSKY, LIZONDO, and REINHART (1998). In our study, we propose an alternative indicator of macroeconomic instability due the recent context of the the euro area's sovereign debt crisis. In fact, the intensification of the financial crisis in September 2008 (in the aftermath of the collapse of Lehman Brothers), has had an serious impact on the EA government bond market and marked the beginning of financial stress for Greece, Ireland, Italy, Portugal, and Spain. As shown in Figure 4.11, after ten years of stability at very low levels, the long-term government bond yields relative to the German Bund have been rising since the beginning of 2010. Due to the unsatisfactory performance of the GIIPS countries group, the spreads was well above those of emerging market countries, such as South Korea and Brazil.

Figure 4.11: Spreads of 10-year government benchmark bonds to German Bund



Source: European Central Bank.

³⁴To obtain real interest rate differential, NOGUEIRA JR. and LEON-LEDESMA (2011) used data on money market interest rates for Mexico and for the United States. CPI inflation was used to obtain the real interest rates from the nominal interest rates collected.

Consequently, we propose to use the sovereign yield spreads to German bonds as an indicator of macroeconomic instability. We expect that this variable would provide some proxy of the risks perceived by foreign firms with respect to the economy under consideration. The widening of sovereign bond yield differentials would indicate the increasing of macroeconomic instability and the loss of confidence in a given economy. In such a case, exporters are willing to shift away from LCP strategy to set prices in their own currencies (PCP strategy), leading to higher extent of ERPT. Using LSTR model, we assume exporter's markup to depend nonlinearly on the importing country's sovereign bond yield differential, that is, when the economy faces a confidence crisis, ERPT becomes higher. The transition variables used as measures of macroeconomic instability in the nonlinear framework is the 10-year government bond yield spreads to the German Bund bys_{t-j} in percentage. The data is obtained from the European Central Bank (ECB) statistics. When the transition variable $s_t = bys_{t-j}$ is exceeding an estimated threshold, these can be interpreted as periods confidence crisis/macroecomic instability. Our smooth transition models are estimated for five heavily indebted EA countries, namely the so-called GIIPS group (Greece, Ireland, Italy, Portugal, and Spain), using monthly data from 1993:01 to 2012:12 in order to cover the changing behaviour in the pass-through dynamics during the EA sovereign debt crisis.

The choice of the adequate lagged bond yield spread bys_{t-j} as a transition variable by means of linearity tests is reported in Table D.10 in Appendix D.3. According to linearity tests, there is a strong evidence of presence of nonlinearities in the five peripheral EA countries. LSTR model is found to be the best specification to capture the nonlinearity with respect to sovereign bond yield differential.

Concerning ERPT estimates, results are reported in Table 4.5. We note that for our 5 GIIPS EA countries we find significant nonlinear response of CPI inflation to exchange rate movements with respect to macroeconomic instability. In the short-run, only Greece and Italy show significant positive relationship between bond yield spread and the extent of pass-through (see Figure D.9 in Appendix D.5). For example, when the Greek bond yield spread (versus Germany) exceeds 2%, the rate of pass-through increases from 0.24% (when $G = 0$) to about 0.42% (when $G = 1$). Otherwise, the nonlinear mechanism is more clear in the long-run. For our sample of 5 EA countries, the extent of ERPT differ strongly in periods of confidence crisis. For Portugal, the ERPT is

equal to 0.32% when yield differential is below 2.14%, but beyond this threshold level, ERPT becomes higher and reaches 0.73% (see Figure 4.12).

Table 4.5: Estimated ERPT elasticities from the LSTR model with $s_t = bys_{t-i}$

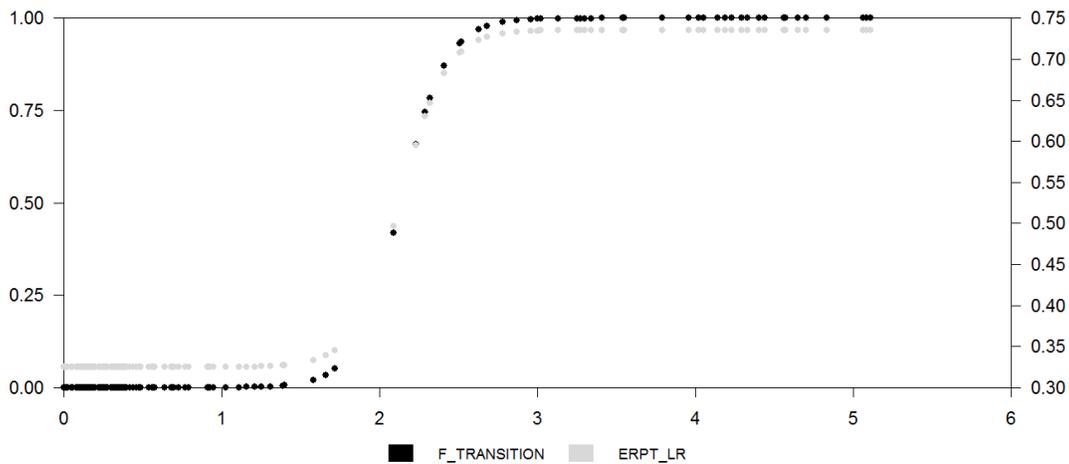
	Greece	Ireland	Italy	Portugal	Spain
Transition variable (s_t)	bys_{t-4}	bys_{t-4}	bys_{t-2}	bys_{t-1}	bys_{t-1}
Threshold (c)	2,720	0,670	2,088	2,137	1,098
	0,000	0,000	0,000	0,000	0,000
Speed of transition (γ)	28,632	14,187	9,084	10,203	20,264
	0,348	0,352	0,326	0,468	0,318
Linear Part : G=0					
<i>SR ERPT</i>	0,243	0,100	0,012	0,163	0,039
	0,004	0,010	0,588	0,000	0,009
<i>LR ERPT</i>	0,325	0,071	0,036	0,325	0,203
	0,002	0,318	0,553	0,003	0,153
Nonlinear Part : G=1					
<i>SR ERPT</i>	0,423	0,382	0,033	0,263	0,106
	0,002	0,281	0,089	0,180	0,191
<i>LR ERPT</i>	0,614	0,782	0,151	0,736	0,472
	0,001	0,140	0,029	0,071	0,093
R^2	0,947	0,788	0,657	0,694	0,737
SSR_{ratio}	0,588	0,676	0,655	0,670	0,796
AIC	-8,531	-8,857	-10,189	-7,267	-8,859
pJB	0,005	0,134	0,628	0,000	0,187
$pLM_{AR(4)}$	0,760	0,922	0,934	0,513	0,439
$pLM_{ARCH(4)}$	0,511	0,878	0,914	0,946	0,184
pLM_C	0,490	0,797	0,198	0,594	0,275
pLM_{RNL}	0,688	0,473	0,363	0,204	0,347

Note: Table reports elasticities of exchange rate pass-through into CPI inflation from LSTR models. Numbers in parentheses are p -values of estimates. R^2 denotes the coefficient of determination, SSR_{ratio} is the ratio of sum of squared residuals between LSTR model and the linear specification, and AIC is the Akaike Information Criterion. The following rows corresponds to the misspecification tests: pJB is the p -values of Jarque-Bera normality test, $pLM_{AR(4)}$ is the p -values of the LM test of no error autocorrelation up to fourth order, $pLM_{ARCH(4)}$ is the p -values of the LM test of no ARCH effects up to fourth order, pLM_C is the p -values of the LM test of parameter constancy and pLM_{RNL} is the p -values of the LM test of no remaining nonlinearity.

Our results suggest that there is an important effect of macroeconomic instability on the ERPT. Under bad economic conditions, firms have no incentive to absorb exchange rate movements in their margins which thus leads to higher ERPT, in opposition with periods of macroeconomic stability when ERPT would be expected to decline. This is in line with NOGUEIRA JR. and LEON-LEDESMA (2011) who found that the sensibility of CPI inflation is higher when Mexican economy faces financial or a confidence crisis. To gain further insight into the role of crisis in determining the degree of pass-through, plots of long-run ERPT estimates over time and past yield spreads are

displayed in Figure 4.13 with the estimated threshold level superimposed. The displayed plots reveal that, since the eruption of the sovereign debt crisis in the beginning of 2010, the transmission of the single currency movements becomes higher after ten years of stability at very low levels. The loss of confidence in GIIPS markets has entailed a higher ERPT rates. This effect might result from foreign firms recognizing that those countries are themselves fundamentally in severe trouble. Indeed, the EA sovereign debt crisis would force exporters to follow PCP strategy due to the general weakness of macroeconomic fundamentals in GIIPS group.

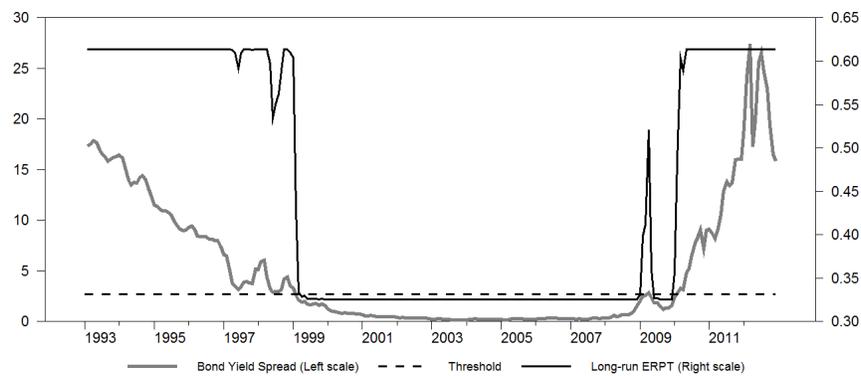
Figure 4.12: Logistic function and long-run ERPT as a function of past yield spread in Portugal



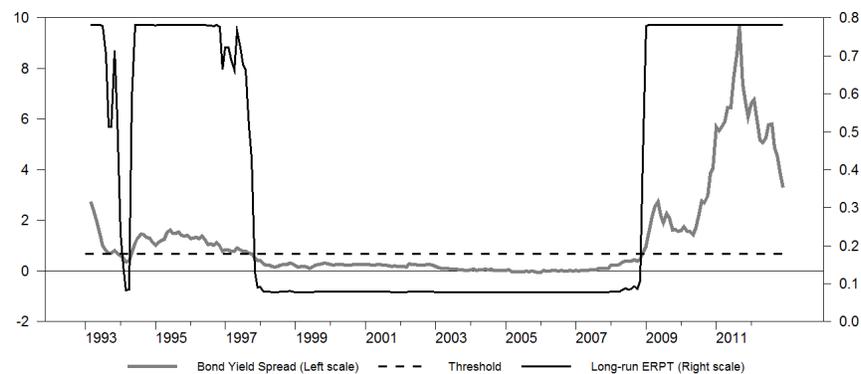
Moreover, a very interesting result is that the 10-year yield spreads versus Germany was very low during the first ten years the third stage of the EMU. During this period, there was a small rate of ERPT throughout our GIIPS EA countries. However, during the pre-EA era, the yield differentials were more pronounced with higher degree of exchange rate transmission. It is plausible that the credibility gained from the adoption of the single currency was responsible for the tightening of bond yield spreads and to some extent to the decline in the rates of ERPT. This conclusion reinforces the argument that the introduction of a set of policies that boost market confidence in the economy can indeed lead to lower ERPT. The adoption of sounder policies may be an effective tool for reducing ERPT. Of course, we do not pretend that all the gain in terms of lower ERPT rates are due to better macroeconomic management or the only source of

nonlinearity, but it is an important finding for the EA countries with historical poor macroeconomic policies. Furthermore, in this context of sovereign debt crisis more attention must be paid to the impact of the euro fluctuations on the CPI inflation. We see that more macroeconomic instability can give rise to higher ERPT, which can be a serious threat to price stability for the Eurozone members. This conclusion has strong policy implications. European monetary authorities must take into account the nonlinear mechanism of ERPT in periods of financial crisis and how it affects inflation dynamics.

Figure 4.13: Time-varying long-run ERPT and past bond yield spread

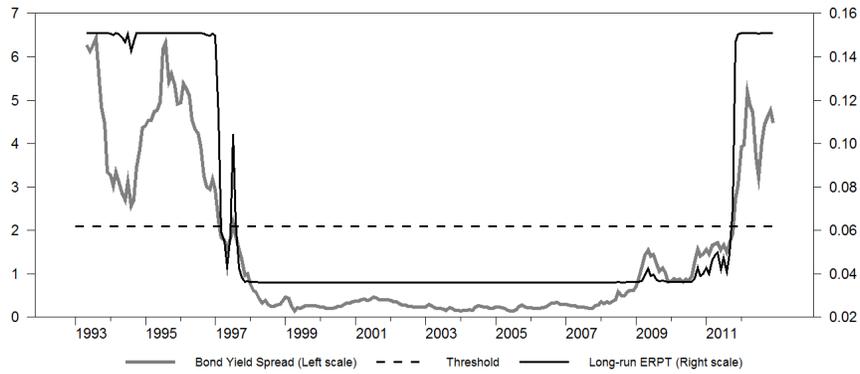


Greece

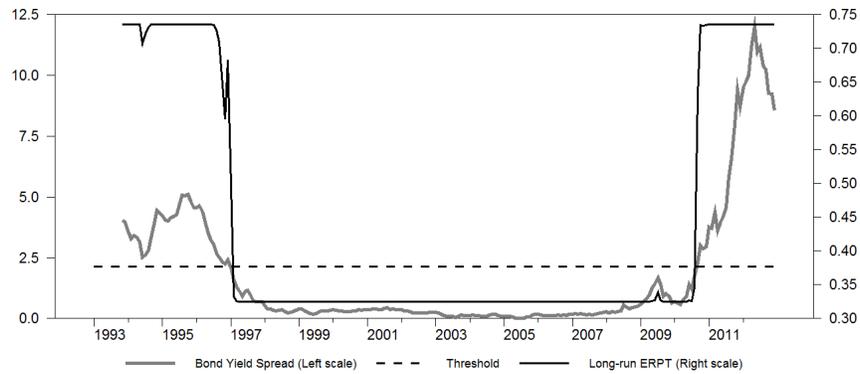


Ireland

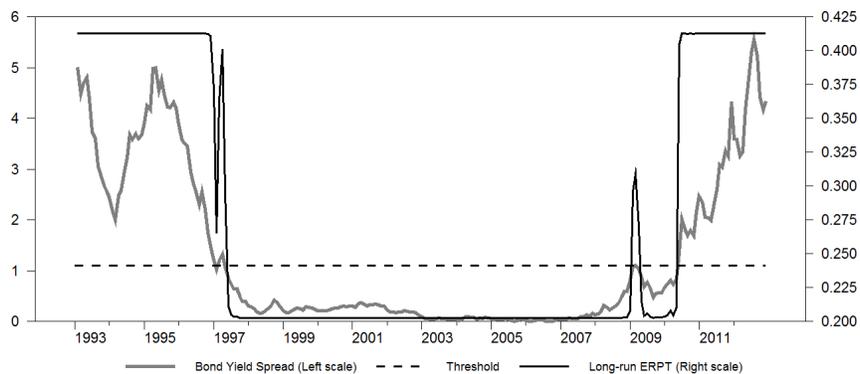
Figure 4.13: Continued



Italy



Portugal



Spain

Note: Time-varying long-run ERPT and past bond yield spread during 1993-2012. Results are from LSTR model with $s_t = bys_{t-i}$.

7. Conclusion

In this study, we investigate for possible nonlinear mechanisms in the exchange rate pass-through (ERPT) to consumer prices for 12 euro area (EA) countries. This exercise is conducted using the family of smooth transition regression models as tool. Mainly, we explore the existence of nonlinearities with respect to three macroeconomic determinants of ERPT, namely inflation environment, exchange rate fluctuations and business cycle.

Using quarterly data spanning from 1975 to 2010, we find strong evidence that pass-through respond non-linearly to inflation level. The transmission of exchange rate is higher when inflation rate surpass some threshold. Results are more striking in the long run with 8 out of 12 EA countries reveal positive relationship between ERPT-Inflation. We give a supportive evidence to the Taylor's view that pass-through is decreasing in a lower and more stable inflation environment. Furthermore, plots of time-varying pass-through coefficients suggest that prices sensibility to exchange rate changes has declined over time in response to a shift to a low-inflation regime.

When considering exchange rate movements as a potential source of nonlinearities, we focus on asymmetries arising from both direction and magnitude of exchange rate. First, we provide a support of asymmetrical ERPT to appreciations and depreciations, but there is no clear-cut about the direction of asymmetry. In other words, for some countries pass-through is found to be greater when exchange rate is depreciating than when it is appreciating. This finding is consistent with the so-called *quantity constraint* theory. Nevertheless, we find the opposite result for the rest of EA countries, i.e. ERPT is higher during importer's currency appreciation than during a period of depreciation. This latter result is line with the *market share* explanation. It is important to note that similar mixed result was pointed out by a number of empirical studies (GIL-PAREJA (2000) and OLIVEI (2002) and COUGHLIN and POLLARD (2004)). Next, we check the asymmetry of pass-through with respect exchange rate magnitude. CPI inflation reaction is found to be higher for large exchange rate changes than for small ones. This can be interpreted as an evidence of the presence of *menu costs*, where large currency movements are promptly transmitted to prices. A careful inspection of time-varying pass-through elasticities reveals that CPI inflation

responsiveness to exchange rate variation was relatively higher during the EMS Crisis and at the launch of the euro.

Thereafter, we consider the business cycle as source of nonlinearities. We report strong evidence that pass-through respond nonlinearly to economic activity in 6 out of 12 EA countries. In other words, the extent of pass-through is found to be different between the periods of expansion and recession in half of EA countries. However, we find no clear direction in this regime-dependence of pass-through to business cycle. In some countries, ERPT is higher during expansions than in recessions; however, in other countries, this result is reversed. These cross-country differences in the nonlinear mechanism of pass-through would have important implications for the design of monetary policy and the control of inflation in the monetary union.

Finally, we test whether periods of macroeconomic instability/confidence crisis may alter the extent of pass-through in a nonlinear way. In the light of the recent European sovereign debt crisis, we propose to use 10-year government bond yield differentials (versus Germany) as an indicator of macroeconomic instability. Our estimation is conducted only for the heavily-indebted EA economies i.e. the GIIPS group (Greece, Ireland, Italy, Portugal, and Spain). Our results show that in periods of widening spreads, which corresponds to episodes of confidence crisis, the degree of ERPT is higher.

Appendix D

D.1. Stationary Tests

Table D.1: Unit Root Tests for π_t

	Austria	Belgium	Germany	Spain	Finland	France
ADF test	-2,49941*	-3,1285**	-2,3269 *	-2,3443*	-2,2732*	-1,3871
ZA test	-4,626*	-5,42664**	-3,28748	-5,20309**	-4,14169*	-4,89641 *
LP test	-5,4525*	-5,7938*	-3,8357	-6,6154*	-4,6757	-4,7807
	Greece	Ireland	Italy	Luxembourg	Netherlands	Portugal
ADF test	-1,0933	-2,4113	-1,5206	-3,1449**	-3,5305**	-1,7559
ZA test	-4,32156*	-4,87005*	-4,71944*	-4,74024**	-4,73330*	-4,94901**
LP test	-4,5474	-5,3954	-6,0446*	-5,0647	-5,8354*	-5,9477*

Key: **, * the null hypothesis of unit root is rejected at 5% and 10% respectively. ZA test (ZIVOT and ANDREWS (1992)) allow for one single break under the alternative hypothesis. LP test (LUMSDAINE and PAPELL (1997)) allow for two structural breaks under the alternative hypothesis. Specifications for ZA and LP tests include both a constant and a time trend. Lag selection: Akaike (AIC). Maximum lags number = 8.

Table D.2: Unit Root Tests for Δe_t

	Austria	Belgium	Germany	Spain	Finland	France
ADF test	-7,07806**	-8,1554 **	-8,4405**	-8,1891**	-8,5382**	-8,1272
ZA test	-8,78526**	-8,64366**	-8,77998**	-8,87720**	-8,98770**	-8,83657**
LP test	-9,8235**	-9,6429**	-9,8282**	-9,7165**	-9,6528**	-9,7718**
	Greece	Ireland	Italy	Luxembourg	Netherlands	Portugal
ADF test	-8,7663	-8,3102	-8,1064**	-8,1554**	-8,4879**	-7,5155**
ZA test	-4,59649*	-8,87136**	-8,81869**	-8,64366**	-8,88402**	-9,07966**
LP test	-5,7138*	-9,6448**	-9,4803**	-9,6429**	-9,9132**	-9,8650**

Key: **, * the null hypothesis of unit root is rejected at 5% and 10% respectively. ZA test (ZIVOT and ANDREWS (1992)) allow for one single break under the alternative hypothesis. LP test (LUMSDAINE and PAPELL (1997)) allow for two structural breaks under the alternative hypothesis. Specifications for ZA and LP tests include both a constant and a time trend. Lag selection: Akaike (AIC). Maximum lags number = 8.

Table D.3: Unit Root Tests for Δw_t^*

	Austria	Belgium	Germany	Spain	Finland	France
ADF test	-6,12781**	-9.1009**	-9.5171**	-9.8040**	-8.6598**	-9.2834**
ZA test	-10.6482**	-3.9223	-10.1216**	-6.51528**	-9.22767**	-6.41160**
LP test	-10.9031**	-4.5495	-10.4572**	-6.7960**	-9.5686**	-6.8802**
	Greece	Ireland	Italy	Luxembourg	Netherlands	Portugal
ADF test	-9.7893**	-10.0655**	-8.7259**	-8.9675**	-9.7030**	-6.8438**
ZA test	-12.4074**	-10.4019**	-9.56370**	-3.91067	-10.0972**	-4.45631
LP test	-12.7675**	-10.4019**	-10.0001**	-4.4399	-10.4022**	-4.6785

Key: **, * the null hypothesis of unit root is rejected at 5% and 10% respectively. ZA test (ZIVOT and ANDREWS (1992)) allow for one single break under the alternative hypothesis. LP test (LUMSDAINE and PAPELL (1997)) allow for two structural breaks under the alternative hypothesis. Specifications for ZA and LP tests include both a constant and a time trend. Lag selection: Akaike (AIC). Maximum lags number = 8.

Table D.4: Unit Root Tests for Δy_t

	Austria	Belgium	Germany	Spain	Finland	France
ADF test	-11,4573**	-6,5366 **	-8,3907**	-3,2332*	-4,2874**	-5,0841**
ZA test	-11,8537**	-7,47145**	-8,50138**	-3,61602	-5,04799*	-5,77605**
LP test	-12,2914**	-7,9376**	-8,7710**	-4,3821	-5,2084	-6,4820*
	Greece	Ireland	Italy	Luxembourg	Netherlands	Portugal
ADF test	-4,1562**	-3,5561**	-6,4883**	-11,2848**	-14,2895**	-4,1707
ZA test	-4,59308*	-4,40607*	-7,51558**	-3,58642	-14,9294**	-5,32930**
LP test	-5,2146	-5,4084	-7,8473**	-4,3088	-15,5851**	-6,1936*

Key: **, * the null hypothesis of unit root is rejected at 5% and 10% respectively. ZA test (ZIVOT and ANDREWS (1992)) allow for one single break under the alternative hypothesis. LP test (LUMSDAINE and PAPELL (1997)) allow for two structural breaks under the alternative hypothesis. Specifications for ZA and LP tests include both a constant and a time trend. Lag selection: Akaike (AIC). Maximum lags number = 8.

Table D.5: Cointegration Tests

	Austria	Belgium	Germany	Spain	Finland	France
EG test	-2,018	-2,724	-2,858	-2,927	-2,039	-2,556
GH test						
<i>constant</i>	-4,407	-4,438	-4,042	-5,021	-4,479	-4,421
<i>constant and slope</i>	-4,883	-5,002	-5,308	-5,768	-6,703*	-5,419
	Greece	Ireland	Italy	Luxembourg	Netherlands	Portugal
EG test	-2,601	-3,337*	-3,414**	-3,257*	-3,313*	-2,786
GH test						
<i>constant</i>	-4,476	-5,191*	-5,439*	-4,806	-5,496*	-4,180
<i>constant and slope</i>	-5,442	-5,454	-5,232	-6,655*	-5,63	-5,704

Key: **, * the null hypothesis of unit root in the residuals (no cointegration) is rejected at 5% and 10% respectively. Specifications for GH test (GREGORY and HANSEN (1996)) include both a constant and a time trend. Lag selection: Akaike (AIC). Maximum lags number = 8.

D.2. Results from linear models

Table D.6: Full results from linear model

	Austria	Belgium	Germany	Spain	Finland	France
Constant	0,003 (0,006)	0,004 (0,000)	0,004 (0,000)	0,001 (0,399)	0,002 (0,106)	0,000 (0,980)
π_{t-1}			0,172 (0,011)			0,355 (0,000)
π_{t-2}	0,174 (0,011)			0,392 (0,000)		
π_{t-3}			0,231 (0,002)			
π_{t-4}	0,514 (0,000)	0,487 (0,000)	0,353 (0,000)	0,458 (0,000)	0,652 (0,000)	0,253 (0,003)
Δe_t	0,040 (0,003)	0,080 (0,000)	0,052 (0,000)	0,124 (0,000)	0,040 (0,011)	0,028 (0,016)
Δe_{t-1}	0,034 (0,012)		0,023 (0,001)		0,046 (0,004)	0,018 (0,147)
Δe_{t-2}		0,042 (0,001)		0,014 (0,358)		
Δe_{t-3}		0,017 (0,035)				0,010 (0,157)
Δe_{t-4}			-0,019 (0,008)	-0,051 (0,026)	0,005 (0,616)	
Δw_t^*	0,075 (0,000)	0,151 (0,000)	0,093 (0,000)	0,202 (0,000)	0,055 (0,034)	0,064 (0,001)
Δw_{t-1}^*	0,069 (0,002)	0,007 (0,622)			0,108 (0,000)	0,028 (0,145)
Δw_{t-2}^*		0,086 (0,000)		-0,061 (0,101)		0,001 (0,015)
Δw_{t-3}^*						
Δw_{t-4}^*						
Δy_t	-0,023 (0,597)	-0,100 (0,181)		0,113 (0,209)	0,079 (0,050)	
Δy_{t-1}			0,043 (0,005)			
Δy_{t-2}	0,068 (0,129)					
Δy_{t-3}			-0,026 (0,133)			
Δy_{t-4}			0,035 (0,074)			
LR ERPT	0,235 (0,000)	0,272 (0,000)	0,115 (0,019)	0,142 (0,003)	0,264 (0,000)	0,142 (0,003)
R^2	0,674	0,666	0,715	0,788	0,734	0,879
SSR	0,002	0,002	0,001	0,007	0,003	0,002
SE of Residuals	0,004	0,004	0,003	0,007	0,005	0,003
AIC	-8,087	-8,068	-8,539	-6,872	-7,570	-8,322
pJB	0,000	0,399	0,000	0,001	0,032	0,000
$pLM_{AR(4)}$	0,819	0,244	0,845	0,209	0,000	0,112
$pLM_{ARCH(4)}$	0,710	0,511	0,869	0,004	0,020	0,860
pRESET	0,051	0,544	0,744	0,657	0,000	0,000

Key: Table reports estimates of linear pass-through equation. Numbers in parentheses are p -values.

Continued

	Greece	Ireland	Italy	Luxembourg	Netherlands	Portugal
Constant	-0.011 (0,000)	-0.001 (0,497)	0.002 (0,208)	0.003 (0,003)	0.000 (0,657)	0.008 (0,014)
π_{t-1}	0.348 (0,000)		0.250 (0,006)			0.138 (0,080)
π_{t-2}		0.284 (0,000)	0.196 (0,024)	0.249 (0,000)	0.129 (0,078)	0.194 (0,014)
π_{t-3}	0.157 (0,034)	0.221 (0,003)	0.222 (0,011)			
π_{t-4}	0.373 (0,000)	0.221 (0,001)	0.205 (0,013)	0.358 (0,000)	0.488 (0,000)	0.252 (0,002)
Δe_t	0.072 (0,006)	0.031 (0,234)	0.060 (0,000)	0.077 (0,000)	0.042 (0,001)	0.104 (0,009)
Δe_{t-1}	0.038 (0,151)	0.120 (0,000)	0.024 (0,151)	0.041 (0,002)		0.076 (0,058)
Δe_{t-2}						
Δe_{t-3}					0.044 (0,000)	
Δe_{t-4}			-0.037 (0,019)			0.093 (0,013)
Δw_t^*	0.118 (0,003)	0.079 (0,057)	0.110 (0,000)	0.145 (0,000)	0.077 (0,000)	0.168 (0,006)
Δw_{t-1}^*	0.063 (0,118)	0.211 (0,000)	0.036 (0,213)	0.044 (0,045)		0.077 (0,226)
Δw_{t-2}^*						
Δw_{t-3}^*					0.069 (0,000)	
Δw_{t-4}^*			-0.049 (0,086)			0.073 (0,254)
Δy_t				-0.024 (0,372)	-0.056 (0,090)	
Δy_{t-1}	0.056 (0,113)		0.182 (0,011)		0.063 (0,042)	
Δy_{t-2}					0.079 (0,008)	
Δy_{t-3}	0.040 (0,274)					
Δy_{t-4}	0.088 (0,010)	0.138 (0,030)	0.050 (0,517)		0.043 (0,150)	0.266 (0,075)
LR ERPT	0.903 (0,012)	0.551 (0,000)	0.447 (0,000)	0.413 (0,000)	0.224 (0,000)	0.657 (0,000)
R^2	0.879	0.712	0.884	0.680	0.682	0.719
SSR	0.011	0.009	0.003	0.002	0.002	0.019
SE of Residuals	0.008	0.008	0.005	0.004	0.004	0.012
AIC	-6.541	-6.641	-7.572	-8.003	-8.270	-5.791
pJB	0.000	0.000	0.000	0.000	0.269	0.000
$pLM_{AR(4)}$	0.560	0.312	0.841	0.491	0.687	0.699
$pLM_{ARCH(4)}$	0.220	0.001	0.009	0.992	0.938	0.000
pRESET	0.012	0.014	0.063	0.202	0.393	0.000

Key: Table reports estimates of linear pass-through equation. Numbers in parentheses are p-values.

D.3. Linearity tests

Table D.7: Linearity tests against STR model with $s_t = \pi_{t-i}$

	Austria				Belgium				Germany			
	π_{t-1}	π_{t-2}	π_{t-3}	π_{t-4}	π_{t-1}	π_{t-2}	π_{t-3}	π_{t-4}	π_{t-1}	π_{t-2}	π_{t-3}	π_{t-4}
H_0	0,455	0,930	0,552	0,013	0,017	0,054	0,000	0,174	0,359	0,549	0,003	0,691
H_{04}	0,588	0,883	0,427	0,019	0,461	0,592	0,123	0,038	0,295	0,394	0,007	0,981
H_{03}	0,285	0,567	0,860	0,262	0,038	0,096	0,514	0,910	0,739	0,866	0,032	0,033
H_{02}	0,238	0,880	0,329	0,229	0,020	0,025	0,000	0,252	0,294	0,433	0,601	0,078
Specification	Linear	Linear	Linear	LSTR	LSTR	Linear	LSTR	Linear	Linear	Linear	LSTR	Linear
	Spain				Finland				France			
	π_{t-1}	π_{t-2}	π_{t-3}	π_{t-4}	π_{t-1}	π_{t-2}	π_{t-3}	π_{t-4}	π_{t-1}	π_{t-2}	π_{t-3}	π_{t-4}
H_0	0,000	0,000	0,000	0,004	0,000	0,000	0,000	0,019	0,000	0,000	0,001	0,000
H_{04}	0,040	0,556	0,001	0,042	0,087	0,028	0,047	0,006	0,000	0,005	0,052	0,243
H_{03}	0,000	0,576	0,011	0,478	0,150	0,002	0,001	0,146	0,020	0,512	0,200	0,004
H_{02}	0,000	0,000	0,010	0,002	0,000	0,000	0,002	0,717	0,012	0,001	0,001	0,001
Specification	ESTR	LSTR	LSTR	LSTR	LSTR	LSTR	ESTR	LSTR	LSTR	LSTR	LSTR	LSTR
	Greece				Ireland				Italy			
	π_{t-1}	π_{t-2}	π_{t-3}	π_{t-4}	π_{t-1}	π_{t-2}	π_{t-3}	π_{t-4}	π_{t-1}	π_{t-2}	π_{t-3}	π_{t-4}
H_0	0,001	0,072	0,020	0,058	0,000	0,000	0,000	0,000	0,000	0,000	0,000	0,001
H_{04}	0,090	0,820	0,016	0,011	0,000	0,002	0,000	0,000	0,001	0,032	0,149	0,061
H_{03}	0,241	0,669	0,642	0,730	0,000	0,000	0,000	0,016	0,000	0,004	0,000	0,060
H_{02}	0,000	0,000	0,057	0,272	0,001	0,000	0,000	0,001	0,000	0,000	0,000	0,008
Specification	LSTR	Linear	LSTR	Linear	LSTR	ESTR	ESTR	LSTR	LSTR	LSTR	LSTR	LSTR
	Luxembourg				Netherlands				Portugal			
	π_{t-1}	π_{t-2}	π_{t-3}	π_{t-4}	π_{t-1}	π_{t-2}	π_{t-3}	π_{t-4}	π_{t-1}	π_{t-2}	π_{t-3}	π_{t-4}
H_0	0,028	0,004	0,256	0,017	0,215	0,011	0,001	0,000	0,003	0,001	0,016	0,036
H_{04}	0,207	0,000	0,501	0,008	0,464	0,349	0,495	0,010	0,058	0,045	0,489	0,228
H_{03}	0,031	0,525	0,193	0,201	0,583	0,025	0,009	0,199	0,018	0,001	0,000	0,138
H_{02}	0,197	0,450	0,286	0,456	0,042	0,025	0,001	0,000	0,000	0,000	0,000	0,000
Specification	ESTR	LSTR	Linear	LSTR	Linear	ESTR	LSTR	LSTR	LSTR	LSTR	LSTR	LSTR

Note: The numbers are p -values of F versions of the LM linearity tests. First row shows the test of linearity against the alternative of STR nonlinearity. The second row until the fourth are the p -values of the sequential test for choosing the adequate transition function. The decision rule is the following: if the test of H_{03} yields the strongest rejection of null hypothesis, we choose the ESTR model. Otherwise, we select the LSTR model. The last row gives the selected specification.

Table D.8: Linearity tests against STR model with $s_t = \Delta e_{t-i}$

Austria												
	Δe_{t-1}	Δe_{t-2}	Δe_{t-3}	Δe_{t-4}	Belgium				Germany			
	Δe_{t-1}	Δe_{t-2}	Δe_{t-3}	Δe_{t-4}	Δe_{t-1}	Δe_{t-2}	Δe_{t-3}	Δe_{t-4}	Δe_{t-1}	Δe_{t-2}	Δe_{t-3}	bf Δe_{t-4}
H_0	0,975	0,116	0,933	0,943	0,001	0,149	0,226	0,162	0,018	0,000	0,956	0,120
H_{04}	0,987	0,311	0,990	0,965	0,014	0,258	0,913	0,028	0,060	0,088	0,978	0,469
H_{03}	0,647	0,088	0,986	0,734	0,605	0,421	0,766	0,763	0,299	0,285	0,917	0,027
H_{01}	0,754	0,329	0,101	0,529	0,001	0,108	0,002	0,476	0,042	0,000	0,298	0,488
Specification	Linear	Linear	Linear	Linear	LSTR	Linear	Linear	Linear	LSTR	LSTR	Linear	Linear
Spain				Finland				France				
	Δe_{t-1}	Δe_{t-2}	Δe_{t-3}	Δe_{t-4}	Δe_{t-1}	Δe_{t-2}	Δe_{t-3}	Δe_{t-4}	Δe_{t-1}	Δe_{t-2}	Δe_{t-3}	bf Δe_{t-4}
H_0	0,028	0,103	0,436	0,206	0,003	0,408	0,981	0,831	0,295	0,439	0,038	0,193
H_{04}	0,036	0,961	0,494	0,492	0,001	0,382	0,986	0,763	0,501	0,703	0,072	0,408
H_{03}	0,115	0,031	0,278	0,439	0,087	0,238	0,663	0,771	0,454	0,205	0,344	0,054
H_{01}	0,390	0,046	0,537	0,065	0,727	0,701	0,796	0,462	0,013	0,071	0,041	0,274
Specification	LSTR	Linear	Linear	Linear	LSTR	Linear	Linear	Linear	Linear	Linear	LSTR	Linear
Greece				Ireland				Italy				
	Δe_{t-1}	Δe_{t-2}	Δe_{t-3}	Δe_{t-4}	Δe_{t-1}	Δe_{t-2}	Δe_{t-3}	Δe_{t-4}	Δe_{t-1}	Δe_{t-2}	Δe_{t-3}	bf Δe_{t-4}
H_0	0,527	0,392	0,600	0,012	0,000	0,000	0,000	0,018	0,004	0,000	0,004	0,119
H_{04}	0,261	0,444	0,239	0,073	0,020	0,060	0,000	0,041	0,057	0,001	0,0661	0,6194
H_{03}	0,796	0,236	0,922	0,042	0,003	0,000	0,319	0,170	0,013	0,8765	0,036	0,087
H_{01}	0,567	0,621	0,565	0,194	0,056	0,421	0,115	0,133	0,196	0,000	0,061	0,105
Specification	Linear	Linear	Linear	ESTR	ESTR	ESTR	LSTR	LSTR	ESTR	LSTR	ESTR	Linear
Luxembourg				Netherlands				Portugal				
	Δe_{t-1}	Δe_{t-2}	Δe_{t-3}	Δe_{t-4}	Δe_{t-1}	Δe_{t-2}	Δe_{t-3}	Δe_{t-4}	Δe_{t-1}	Δe_{t-2}	Δe_{t-3}	bf Δe_{t-4}
H_0	0,010	0,062	0,618	0,463	0,177	0,124	0,095	0,037	0,012	0,011	0,926	0,908
H_{04}	0,222	0,417	0,877	0,198	0,336	0,780	0,090	0,129	0,192	0,032	0,900	0,842
H_{03}	0,098	0,121	0,306	0,497	0,198	0,160	0,459	0,384	0,050	0,076	0,790	0,948
H_{01}	0,012	0,056	0,372	0,778	0,271	0,018	0,182	0,028	0,041	0,251	0,544	0,322
Specification	LSTR	Linear	Linear	Linear	Linear	Linear	Linear	LSTR	LSTR	LSTR	Linear	Linear

Note: The numbers are p -values of F versions of the LM linearity tests. First row shows the test of linearity against the alternative of STR nonlinearity. The second row until the fourth row shows the p -values of the sequential test for choosing the adequate transition function. The decision rule is the following: if the test of H_{03} yields the strongest rejection of null hypothesis, we choose the ESTR model. Otherwise, we select the LSTR model. The last row gives the selected specification.

Table D.9: Linearity tests against STR model with $s_t = \Delta y_{t-i}$

	Austria				Belgium				Germany			
	Δy_{t-1}	Δy_{t-2}	Δy_{t-3}	Δy_{t-4}	Δy_{t-1}	Δy_{t-2}	Δy_{t-3}	Δy_{t-4}	Δy_{t-1}	Δy_{t-2}	Δy_{t-3}	Δy_{t-4}
H_0	0,183	0,933	0,009	0,035	0,010	0,837	0,040	0,349	0,373	0,032	0,011	0,042
H_{04}	0,056	0,986	0,016	0,054	0,128	0,666	0,025	0,373	0,162	0,278	0,023	0,212
H_{03}	0,991	0,100	0,155	0,351	0,001	0,818	0,679	0,176	0,602	0,007	0,543	0,082
H_{01}	0,519	0,823	0,281	0,102	0,083	0,813	0,388	0,829	0,581	0,475	0,023	0,137
Specification	Linear	Linear	LSTR	LSTR	ESTR	Linear	LSTR	Linear	Linear	ESTR	LSTR	ESTR
	Spain				Finland				France			
	Δy_{t-1}	Δy_{t-2}	Δy_{t-3}	Δy_{t-4}	Δy_{t-1}	Δy_{t-2}	Δy_{t-3}	Δy_{t-4}	Δy_{t-1}	Δy_{t-2}	Δy_{t-3}	Δy_{t-4}
H_0	0,339	0,453	0,044	0,473	0,319	0,039	0,039	0,037	0,178	0,593	0,136	0,144
H_{04}	0,292	0,811	0,531	0,634	0,701	0,030	0,035	0,139	0,180	0,684	0,001	0,195
H_{03}	0,322	0,078	0,007	0,146	0,221	0,412	0,696	0,809	0,486	0,308	0,800	0,589
H_{01}	0,649	0,534	0,165	0,691	0,201	0,169	0,053	0,005	0,199	0,576	0,019	0,085
Specification	Linear	Linear	ESTR	Linear	Linear	LSTR	LSTR	LSTR	Linear	Linear	Linear	Linear
	Greece				Ireland				Italy			
	Δy_{t-1}	Δy_{t-2}	Δy_{t-3}	Δy_{t-4}	Δy_{t-1}	Δy_{t-2}	Δy_{t-3}	Δy_{t-4}	Δy_{t-1}	Δy_{t-2}	Δy_{t-3}	Δy_{t-4}
H_0	0,001	0,000	0,000	0,012	0,373	0,304	0,947	0,403	0,000	0,000	0,000	0,080
H_{04}	0,798	0,000	0,047	0,139	0,857	0,894	0,921	0,036	0,056	0,102	0,280	0,267
H_{03}	0,000	0,017	0,000	0,018	0,175	0,050	0,789	0,971	0,000	0,000	0,000	0,416
H_{01}	0,093	0,248	0,064	0,176	0,095	0,166	0,571	0,878	0,000	0,005	0,000	0,032
Specification	ESTR	LSTR	ESTR	ESTR	Linear	Linear	Linear	Linear	ESTR	ESTR	ESTR	Linear
	Luxembourg				Netherlands				Portugal			
	Δy_{t-1}	Δy_{t-2}	Δy_{t-3}	Δy_{t-4}	Δy_{t-1}	Δy_{t-2}	Δy_{t-3}	Δy_{t-4}	Δy_{t-1}	Δy_{t-2}	Δy_{t-3}	Δy_{t-4}
H_0	0,785	0,473	0,978	0,360	0,017	0,006	0,047	0,025	0,669	0,025	0,033	0,003
H_{04}	0,964	0,510	0,837	0,716	0,009	0,004	0,148	0,066	0,897	0,282	0,192	0,373
H_{03}	0,852	0,537	0,867	0,090	0,249	0,260	0,380	0,045	0,674	0,055	0,031	0,000
H_{01}	0,070	0,295	0,884	0,512	0,322	0,171	0,037	0,410	0,038	0,017	0,200	0,229
Specification	Linear	Linear	Linear	Linear	LSTR	LSTR	LSTR	ESTR	Linear	LSTR	ESTR	ESTR

Note: The numbers are p -values of F versions of the LM linearity tests. First row shows the test of linearity against the alternative of STR nonlinearity. The second row until the fourth are the p -values of the sequential test for choosing the adequate transition function. The decision rule is the following: if the test of H_{03} yields the strongest rejection of null hypothesis, we choose the ESTR model. Otherwise, we select the LSTR model. The last row gives the selected specification.

Table D.10: Linearity tests against STR model with $s_t = bys_{t-j}$

Country	Transition Variable	H_0	H_{04}	H_{03}	H_{01}	Specification
Greece	bys_{t-1}	0,026	0,720	0,139	0,000	LSTR
	bys_{t-2}	0,128	0,738	0,386	0,000	Linear
	bys_{t-3}	0,079	0,685	0,296	0,000	Linear
	bys_{t-4}	0,047	0,080	0,210	0,256	LSTR
Ireland	bys_{t-1}	0,058	0,185	0,265	0,063	Linear
	bys_{t-2}	0,009	0,121	0,073	0,026	LSTR
	bys_{t-3}	0,122	0,678	0,198	0,014	Linear
	bys_{t-4}	0,002	0,571	0,143	0,000	LSTR
Italy	bys_{t-1}	0,000	0,042	0,002	0,040	ESTR
	bys_{t-2}	0,000	0,001	0,013	0,023	LSTR
	bys_{t-3}	0,000	0,032	0,006	0,011	ESTR
	bys_{t-4}	0,000	0,057	0,007	0,005	LSTR
Portugal	bys_{t-1}	0,002	0,017	0,081	0,077	LSTR
	bys_{t-2}	0,058	0,185	0,265	0,063	Linear
	bys_{t-3}	0,026	0,035	0,563	0,047	LSTR
	bys_{t-4}	0,031	0,035	0,803	0,018	LSTR
Spain	bys_{t-1}	0,003	0,049	0,092	0,015	LSTR
	bys_{t-2}	0,003	0,043	0,165	0,006	LSTR
	bys_{t-3}	0,004	0,077	0,177	0,003	LSTR
	bys_{t-4}	0,006	0,112	0,180	0,003	LSTR

Note: The numbers are p -values of F versions of the LM linearity tests. First row shows the test of linearity against the alternative of STR nonlinearity. The second row until the forth are the p -values of the sequential test for choosing the adequate transition function. The decision rule is the following: if the test of H_{03} yields the strongest rejection of null hypothesis, we choose the ESTR model. Otherwise, we select the LSTR model. The last row gives the selected model.

D.4. Full results from STR pass-through models

Table D.11: Estimation results from LSTR model with $s_t = \pi_{t-i}$

	Austria	Belgium	Germany	Spain	Finland	France
s_t	π_{t-4}	π_{t-1}	π_{t-1}	π_{t-4}	π_{t-3}	π_{t-2}
c	0,033 (0,000)	0,030 (0,000)	0,013 (0,000)	0,022 (0,000)	0,027 (0,000)	0,011 (0,000)
γ	22,013 (0,547)	17,566 (0,312)	9,390 (0,208)	12,702 (0,437)	13,291 (0,531)	6,134 (0,067)
Linear Part: $G = 0$						
Constant	0,000 (0,866)	0,007 (0,000)	0,005 (0,000)	0,001 (0,753)	0,001 (0,485)	0,003 (0,021)
π_{t-1}			0,195 (0,058)			0,174 (0,262)
π_{t-2}	0,160 (0,077)					
π_{t-3}				-0,115 (0,287)		
π_{t-4}	0,534 (0,534)	0,068 (0,638)	0,438 (0,000)	0,863 (0,000)	0,782 (0,000)	0,257 (0,053)
Δe_t	0,043 (0,042)	0,091 (0,000)	0,063 (0,000)	0,085 (0,009)	0,044 (0,005)	0,066 (0,001)
Δe_{t-1}	0,004 (0,821)	0,042 (0,033)			0,048 (0,002)	-0,013 (0,535)
Δe_{t-2}				0,015 (0,626)		
Δe_{t-3}		-0,002 (0,845)				-0,004 (0,680)
Δe_{t-4}			-0,021 (0,116)	0,010 (0,567)	-0,003 (0,785)	
Δw_t^*	0,078 (0,024)	0,140 (0,000)	0,091 (0,000)	0,171 (0,003)	0,058 (0,020)	0,111 (0,002)
Δw_{t-1}^*	0,047 (0,213)	0,004 (0,844)			0,098 (0,000)	-0,012 (0,723)
Δw_{t-2}^*		0,087 (0,011)	-0,001 (0,956)	0,040 (0,482)		
Δw_{t-3}^*						
Δw_{t-4}^*			-0,006 (0,788)			
Δy_t	0,117 (0,206)	-0,293 (0,004)		0,065 (0,534)	0,142 (0,000)	0,000 (0,517)
Δy_{t-1}			0,026 (0,148)			
Δy_{t-2}	0,036 (0,708)					
Δy_{t-3}			-0,036 (0,085)			
Δy_{t-4}			0,063 (0,013)			
Non-linear Part: $G = 1$						
Δe_t	-0,018 (0,516)	-0,017 (0,529)	-0,060 (0,489)	0,082 (0,063)	-0,024 (0,753)	-0,071 (0,015)
Δe_{t-1}	0,071 (0,014)				-0,016 (0,762)	0,080 (0,004)
Δe_{t-2}		-0,014 (0,595)		-0,097 (0,026)		
Δe_{t-3}		0,044 (0,011)				0,046 (0,014)
Δe_{t-4}			0,110 (0,245)	0,057 (0,078)	0,120 (0,079)	

Key: Table reports estimates of STR pass-through equation. Numbers in parentheses are p -values.

Continued

	Greece	Ireland	Italy	Luxembourg	Netherlands	Portugal
s_t	π_{t-3}	π_{t-2}	π_{t-2}	π_{t-4}	π_{t-3}	π_{t-1}
c	0,022 (0,024)	0,034 (0,000)	0,031 (0,000)	0,015 (0,000)	0,008 (0,000)	0,088 (0,000)
γ	2,358 (0,120)	8,456 (0,003)	2,449 (0,002)	4,909 (0,056)	9,361 (0,333)	4,061 (0,053)
Linear Part: $G = 0$						
Constant	-0,014 (0,346)	-0,001 (0,551)	0,002 (0,106)	0,002 (0,048)	-0,002 (0,161)	0,002 (0,716)
π_{t-1}			0,552 (0,000)			0,013 (0,965)
π_{t-2}		0,453 (0,000)		0,249 (0,007)	0,064 (0,571)	0,216 (0,443)
π_{t-3}		0,092 (0,250)	-0,067 (0,635)			-0,119 (0,702)
π_{t-4}	0,645 (0,000)	0,189 (0,015)	0,220 (0,096)	0,541 (0,000)	0,490 (0,000)	
Δe_t	0,105 (0,134)	0,043 (0,097)	0,032 (0,050)	0,053 (0,002)	0,049 (0,009)	0,040 (0,547)
Δe_{t-1}	-0,046 (0,541)	0,074 (0,003)	0,021 (0,221)	0,039 (0,012)		0,001 (0,993)
Δe_{t-2}						
Δe_{t-3}					0,010 (0,327)	
Δe_{t-4}						0,012 (0,765)
Δw_t^*	0,187 (0,036)	0,113 (0,005)	0,074 (0,007)	0,098 (0,001)	0,058 (0,072)	0,099 (0,300)
Δw_{t-1}^*	-0,029 (0,765)	0,139 (0,000)	0,025 (0,429)	0,059 (0,020)		0,011 (0,910)
Δw_{t-2}^*						
Δw_{t-3}^*					0,001 (0,944)	
Δw_{t-4}^*						
Δy_t	-0,037 (0,728)			-0,007 (0,808)	-0,028 (0,679)	
Δy_{t-1}			0,038 (0,616)		0,043 (0,501)	
Δy_{t-2}					0,115 (0,020)	
Δy_{t-3}		0,052 (0,298)	0,027 (0,755)			
Δy_{t-4}	-0,068 (0,544)	0,127 (0,019)			0,099 (0,099)	0,056 (0,787)
Non-linear Part: $G = 1$						
Δe_t	-0,049 (0,595)	1,871 (0,001)	0,134 (0,013)	0,106 (0,008)	-0,013 (0,647)	0,045 (0,609)
Δe_{t-1}	0,191 (0,062)	-1,355 (0,018)	0,078 (0,191)	0,023 (0,531)		0,165 (0,076)
Δe_{t-2}						
Δe_{t-3}					0,034 (0,062)	
Δe_{t-4}						0,175 (0,005)

Key: Table reports estimates of STR pass-through equation. Numbers in parentheses are p -values.

Table D.12: Estimation results from LSTR model with $s_t = \Delta e_{t-i}$

	Belgique	Grèce	Italie	Luxembourg	Portugal
s_t	Δe_{t-4}	Δe_{t-4}	Δe_{t-2}	Δe_{t-1}	Δe_{t-1}
c	0,004 (0,050)	-0,021 (0,000)	0,044 (0,000)	0,037 (0,000)	0,045 (0,000)
γ	60,750 (0,555)	9,675 (0,262)	7,513 (0,095)	18,530 (0,379)	5,317 (0,029)
Linear part: G=0					
Constant	0,005 (0,001)	-0,008 (0,331)	0,000 (0,925)	0,003 (0,009)	0,002 (0,622)
π_{t-1}		0,545 (0,090)	0,478 (0,000)		
π_{t-2}				0,211 (0,006)	
π_{t-3}		-0,187 (0,484)	0,089 (0,360)		
π_{t-4}	0,445 (0,000)	0,439 (0,071)	0,227 (0,013)	0,330 (0,000)	0,662 (0,000)
Δe_t	0,101 (0,000)	0,196 (0,033)	0,037 (0,030)	0,060 (0,000)	0,069 (0,131)
Δe_{t-1}		-0,091 (0,249)	0,052 (0,008)	0,020 (0,275)	
Δe_{t-2}	0,032 (0,083)				
Δe_{t-3}	0,024 (0,086)				-0,035 (0,466)
Δe_{t-4}					
Δw_t^*	0,193 (0,000)	0,255 (0,014)	0,075 (0,005)	0,115 (0,000)	0,087 (0,240)
Δw_{t-1}^*	0,022 (0,196)	-0,114 (0,314)	0,098 (0,005)	0,049 (0,057)	0,013 (0,816)
Δw_{t-2}^*	0,080 (0,002)				
Δw_{t-3}^*					-0,066 (0,323)
Δw_{t-4}^*					
Δy_t	-0,167 (0,129)			-0,050 (0,075)	
Δy_{t-1}		-0,119 (0,339)	0,321 (0,000)		
Δy_{t-2}					
Δy_{t-3}		0,146 (0,163)	0,007 (0,931)		0,201 (0,334)
Δy_{t-4}		0,118 (0,139)			0,026 (0,904)
Non-linear part: G=1					
Δe_t	-0,060 (0,029)	-0,147 (0,131)	0,064 (0,330)	0,063 (0,109)	0,203 (0,010)
Δe_{t-1}		0,132 (0,129)	-0,175 (0,004)	-0,051 (0,337)	
Δe_{t-2}	-0,005 (0,849)				
Δe_{t-3}	-0,009 (0,594)				0,448 (0,000)
Δe_{t-4}					

Key: Table reports estimates of LSTR pass-through equation. Numbers in parentheses are p -values.

Table D.13: Estimation results from ESTR model with $s_t = \Delta e_{t-i}$

	Belgique	Allemagne	Espagne	Finlande	France	Grèce	Irlande	Italie	Luxembourg	Pays-Bas
s_t	Δe_{t-4}	Δe_{t-1}	Δe_{t-4}	Δe_{t-2}	Δe_{t-3}	Δe_{t-3}	Δe_{t-2}	Δe_{t-1}	Δe_{t-3}	Δe_{t-4}
c	0,022 (0,059)	0,006 (0,037)	0,035 (0,004)	0,021 (0,000)	-0,022 (0,000)	0,030 (0,000)	0,043 (0,000)	0,016 (0,000)	0,010 (0,016)	0,033 (0,000)
γ	4,381 (0,000)	11,092 (0,062)	4,322 (0,110)	11,347 (0,004)	2,487 (0,064)	33,264 (0,053)	1,274 (0,025)	9,112 (0,105)	4,041 (0,057)	1,128 (0,058)
Linear Part: G=0										
Constant	0,004 (0,188)	0,005 (0,181)	-0,005 (0,306)	0,007 (0,166)	-0,002 (0,332)	0,002 (0,956)	0,003 (0,504)	0,003 (0,532)	0,009 (0,001)	-0,003 (0,226)
π_{t-1}			0,028 (0,896)			-0,495 (0,591)	0,160 (0,267)	0,812 (0,015)		0,391 (0,032)
π_{t-2}		0,798 (0,007)					-0,169 (0,301)		0,174 (0,478)	
π_{t-3}					0,410 (0,007)	0,049 (0,919)		-0,429 (0,274)		
π_{t-4}	0,696 (0,000)		0,748 (0,003)	0,139 (0,459)	0,345 (0,033)	2,068 (0,044)	0,304 (0,014)	0,410 (0,100)	0,290 (0,090)	0,506 (0,000)
Δe_t	-0,016 (0,681)	0,002 (0,972)	0,019 (0,814)	-0,071 (0,183)	-0,019 (0,485)	-0,291 (0,073)	0,065 (0,256)	0,009 (0,886)	0,055 (0,062)	-0,018 (0,476)
Δe_{t-1}				-0,046 (0,691)	0,065 (0,015)	0,034 (0,852)	0,023 (0,652)	0,014 (0,943)	-0,006 (0,878)	0,021 (0,162)
Δe_{t-2}	-0,002 (0,967)		-0,065 (0,128)							
Δe_{t-3}	-0,015 (0,449)	-0,099 (0,015)								
Δe_{t-4}			-0,019 (0,611)	0,006 (0,864)	-0,038 (0,027)			-0,067 (0,125)		
Δw_t^*	-0,064 (0,479)	-0,029 (0,779)	0,059 (0,731)	-0,322 (0,029)	0,045 (0,159)	-0,455 (0,006)	0,062 (0,594)	0,067 (0,426)	0,042 (0,393)	0,017 (0,727)
Δw_{t-1}^*	-0,041 (0,370)		-0,061 (0,449)	0,012 (0,932)	0,112 (0,005)	0,186 (0,586)	0,176 (0,026)	-0,117 (0,279)	0,016 (0,820)	
Δw_{t-2}^*	0,022 (0,786)	-0,129 (0,008)								
Δw_{t-3}^*										0,046 (0,024)
Δw_{t-4}^*		0,009 (0,886)								
Δy_t	-0,111 (0,585)		0,012 (0,951)	-0,163 (0,475)					-0,174 (0,078)	0,113 (0,234)
Δy_{t-1}						-0,258 (0,412)		0,047 (0,817)		0,195 (0,003)
Δy_{t-2}		0,165 (0,062)			0,338 (0,024)					-0,015 (0,826)
Δy_{t-3}						0,083 (0,808)	0,037 (0,745)	0,084 (0,784)		
Δy_{t-4}		0,100 (0,246)				0,578 (0,175)	0,080 (0,567)			0,044 (0,416)
Non-linear Part: G=1										
Δe_t	0,119 (0,006)	0,073 (0,200)	0,102 (0,253)	0,121 (0,040)	0,095 (0,008)	0,395 (0,018)	-0,076 (0,294)	0,061 (0,370)	0,035 (0,329)	0,122 (0,002)
Δe_{t-1}				0,101 (0,405)	-0,059 (0,110)	0,009 (0,961)	0,140 (0,042)	0,022 (0,912)	0,058 (0,234)	-0,002 (0,920)
Δe_{t-2}	0,045 (0,333)		0,054 (0,259)							
Δe_{t-3}	0,042 (0,076)	0,112 (0,009)								
Δe_{t-4}			0,059 (0,160)	-0,006 (0,871)	0,046 (0,038)			0,057 (0,227)		

Key: Table reports estimates of ESTR pass-through equation. Numbers in parentheses are p -values.

Table D.14: Estimation results from LSTR model with $s_t = \Delta y_{t-i}$

	Autriche	Belgique	Allemagne	Espagne	Finlande	Greece	Italy	Netherlands	Portugal
s_t	Δy_{t-1}	Δy_{t-3}	Δy_{t-4}	Δy_{t-3}	Δy_{t-2}	Δy_{t-2}	Δy_{t-1}	Δy_{t-4}	Δy_{t-3}
c	0,040 (0,000)	0,003 (0,000)	0,010 (0,079)	0,006 (0,509)	0,029 (0,000)	0,021 (0,009)	0,017 (0,000)	0,007 (0,000)	0,013 (0,000)
γ	24,444 (0,651)	20,760 (0,168)	3,304 (0,162)	26,210 (0,000)	3,740 (0,193)	4,585 (0,202)	3,944 (0,003)	8,959 (0,265)	26,378 (0,311)
Linear Part: G=0									
Constant	0,002 (0,193)	0,009 (0,000)	0,007 (0,000)	0,000 (0,960)	0,000 (0,931)	0,001 (0,603)	0,001 (0,094)	-0,001 (0,649)	0,006 (0,091)
π_{t-1}			0,352 (0,004)				0,388 (0,000)		0,151 (0,076)
π_{t-2}	0,197 (0,007)			0,091 (0,553)				0,076 (0,490)	0,293 (0,003)
π_{t-3}							0,233 (0,000)		
π_{t-4}	0,538 (0,000)	0,425 (0,000)	0,167 (0,273)	0,765 (0,000)	0,673 (0,000)	0,728 (0,000)	0,206 (0,001)	0,478 (0,000)	0,308 (0,000)
Δe_t	0,044 (0,001)	0,105 (0,000)	0,024 (0,269)	0,049 (0,129)	0,010 (0,708)	0,112 (0,001)	0,044 (0,000)	0,043 (0,025)	0,093 (0,021)
Δe_{t-1}				0,056 (0,032)	0,031 (0,234)	0,046 (0,294)	0,013 (0,334)		
Δe_{t-2}		0,058 (0,007)		-0,077 (0,023)					0,041 (0,146)
Δe_{t-3}		0,025 (0,044)						0,050 (0,014)	
Δe_{t-4}	0,006 (0,685)		0,019 (0,477)		0,008 (0,621)				0,041 (0,129)
Δw_t^*	0,084 (0,000)	0,168 (0,012)	0,028 (0,445)	0,125 (0,002)	-0,018 (0,719)	0,186 (0,000)	0,103 (0,000)	0,029 (0,439)	0,155 (0,016)
Δw_{t-1}^*	0,016 (0,256)	-0,067 (0,090)			0,102 (0,014)	0,101 (0,100)	0,008 (0,729)		0,000 (0,995)
Δw_{t-2}^*		0,171 (0,000)	0,039 (0,154)	-0,046 (0,376)					
Δw_{t-3}^*								0,086 (0,004)	
Δw_{t-4}^*	-0,001 (0,982)		0,051 (0,279)						
Δy_t	0,027 (0,642)	-0,389 (0,016)		0,565 (0,001)	-0,013 (0,864)	-0,025 (0,605)		-0,056 (0,232)	
Δy_{t-1}			0,006 (0,839)				0,042 (0,528)	0,020 (0,688)	
Δy_{t-2}	0,079 (0,085)							0,041 (0,422)	
Δy_{t-3}			-0,041 (0,145)				-0,011 (0,848)		0,014 (0,957)
Δy_{t-4}			-0,078 (0,328)			0,059 (0,203)		-0,066 (0,275)	0,370 (0,077)
Non-linear Part: G=1									
Δe_t	0,178 (0,046)	0,057 (0,645)	0,112 (0,052)	0,114 (0,010)	0,070 (0,126)	-0,105 (0,258)	0,029 (0,895)	-0,011 (0,684)	0,032 (0,743)
Δe_{t-1}				-0,060 (0,066)	0,023 (0,582)	0,023 (0,774)	-0,321 (0,213)		
Δe_{t-2}		-0,034 (0,300)		0,069 (0,151)					0,080 (0,299)
Δe_{t-3}		-0,029 (0,275)						-0,012 (0,660)	
Δe_{t-4}	-0,139 (0,042)		-0,068 (0,232)		0,012 (0,685)				0,114 (0,159)

Key: Table reports estimates of LSTR pass-through equation. Numbers in parentheses are p-values.

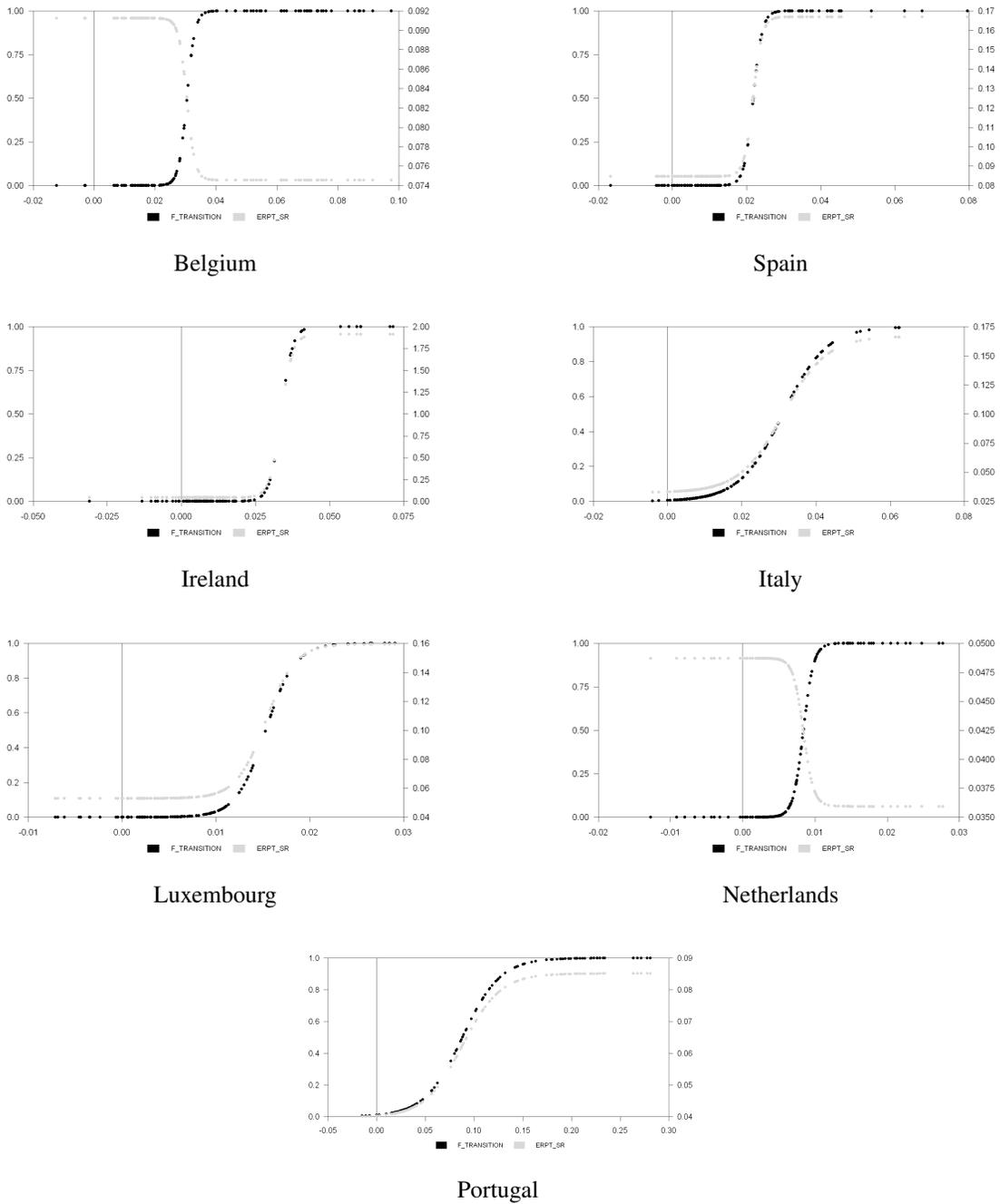
Table D.15: Estimation results from LSTR model with $s_t = by_{s_{t-i}}$

	Greece	Ireland	Italy	Portugal	Spain
Transition variable (s_t)	$by_{s_{t-3}}$	$by_{s_{t-2}}$	$by_{s_{t-2}}$	$by_{s_{t-1}}$	$by_{s_{t-4}}$
Threshold (c)	2,720	0,670	2,088	2,137	1,098
	0,000	0,000	0,000	0,000	0,000
Speed of transition (γ)	28,632	14,187	9,084	10,203	20,264
	0,348	0,352	0,326	0,468	0,318
Linear Part: G=0					
Constant	-0,005	-0,006	0,002	0,006	-0,002
	0,043	0,000	0,000	0,002	0,005
π_{t-1}	-0,401	0,150		0,366	0,287
	0,000	0,144		0,000	0,001
π_{t-2}			0,143		
			0,109		
π_{t-3}		0,147			
		0,141			
π_{t-4}	0,120		0,283		0,197
	0,231		0,003		0,019
Δe_t	0,243	0,100	0,012	0,163	0,039
	0,004	0,010	0,588	0,000	0,009
Δe_{t-1}			0,041		
			0,074		
Δe_{t-2}	0,174				0,022
	0,037				0,704
Δe_{t-3}			-0,033		
			0,153		
Δe_{t-4}		-0,045		0,043	
		0,207		0,320	
Δw_t^*	0,059	0,070	0,030	0,375	0,024
	0,014	0,005	0,006	0,000	0,223
Δw_{t-1}^*	0,040		0,020	0,200	
	0,125		0,064	0,005	
Δw_{t-2}^*	0,062		0,011		
	0,014		0,312		
Δw_{t-3}^*		-0,045		-0,115	0,024
		0,125		0,112	0,275
Δw_{t-4}^*			-0,014		
			0,194		
Δy_t				-0,030	
				0,166	
Δy_{t-1}					0,012
					0,445
Δy_{t-2}		-0,007	0,009		0,019
		0,226	0,261		0,208
Δy_{t-3}	-0,023				
	0,188				
Δy_{t-4}	-0,026	0,006			
	0,072	0,232			
Nonlinear Part: G=1					
Δe_t	0,189	0,283	0,021	0,100	0,068
	0,103	0,429	0,485	0,615	0,387
Δe_{t-1}		0,213	-0,039		0,041
		0,120	0,233		0,412
Δe_{t-2}	0,180			-0,138	0,029
	0,101			0,118	0,572
Δe_{t-3}			0,045		
			0,139		
Δe_{t-4}			0,039	0,299	
			0,184	0,014	

Key: Table reports estimates of LSTR pass-through equation. Numbers in parentheses are p-values.

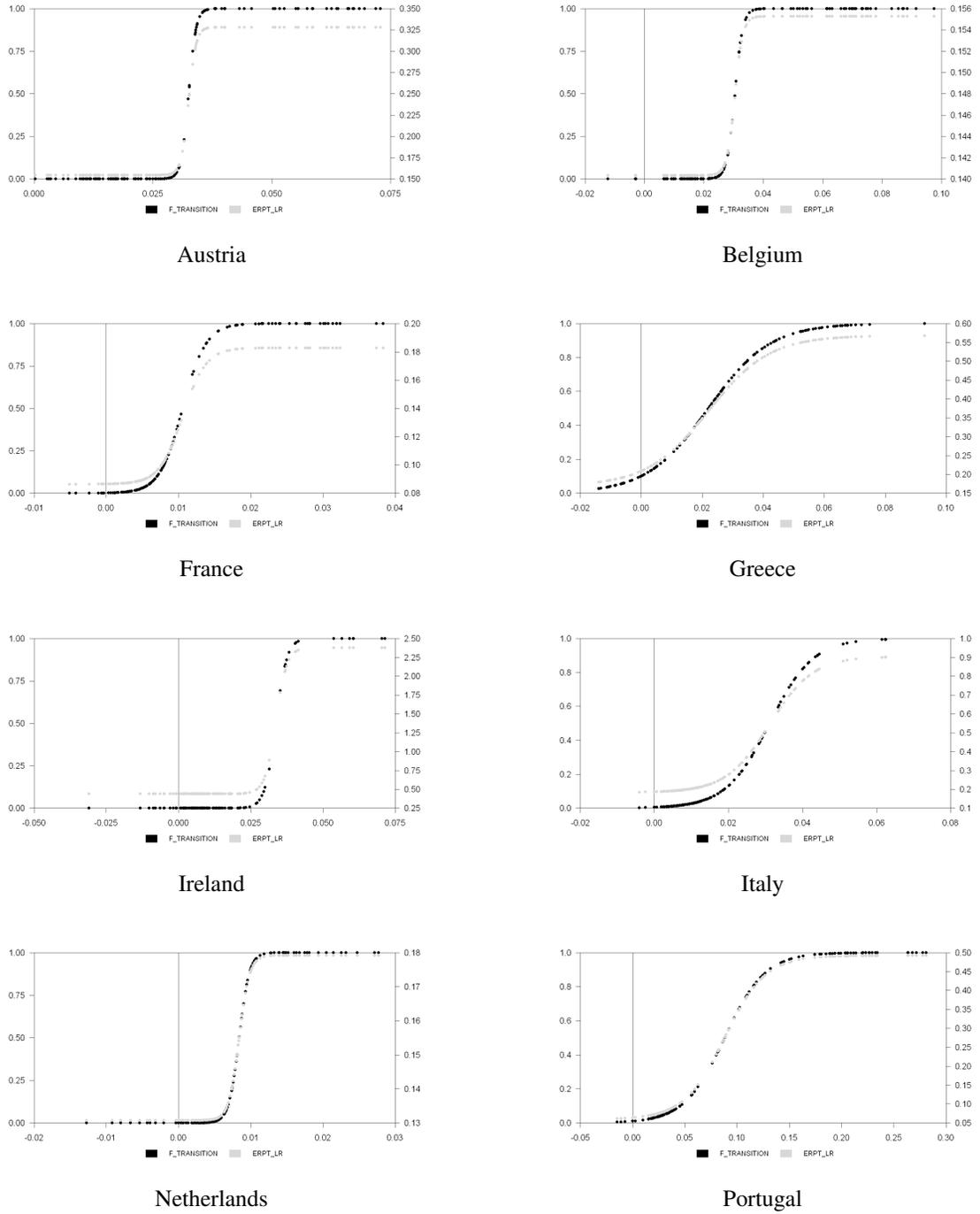
D.5. Plots of estimated transition function and ERPT

Figure D.1: Logistic functions and short-run ERPT as a function of past inflation rates



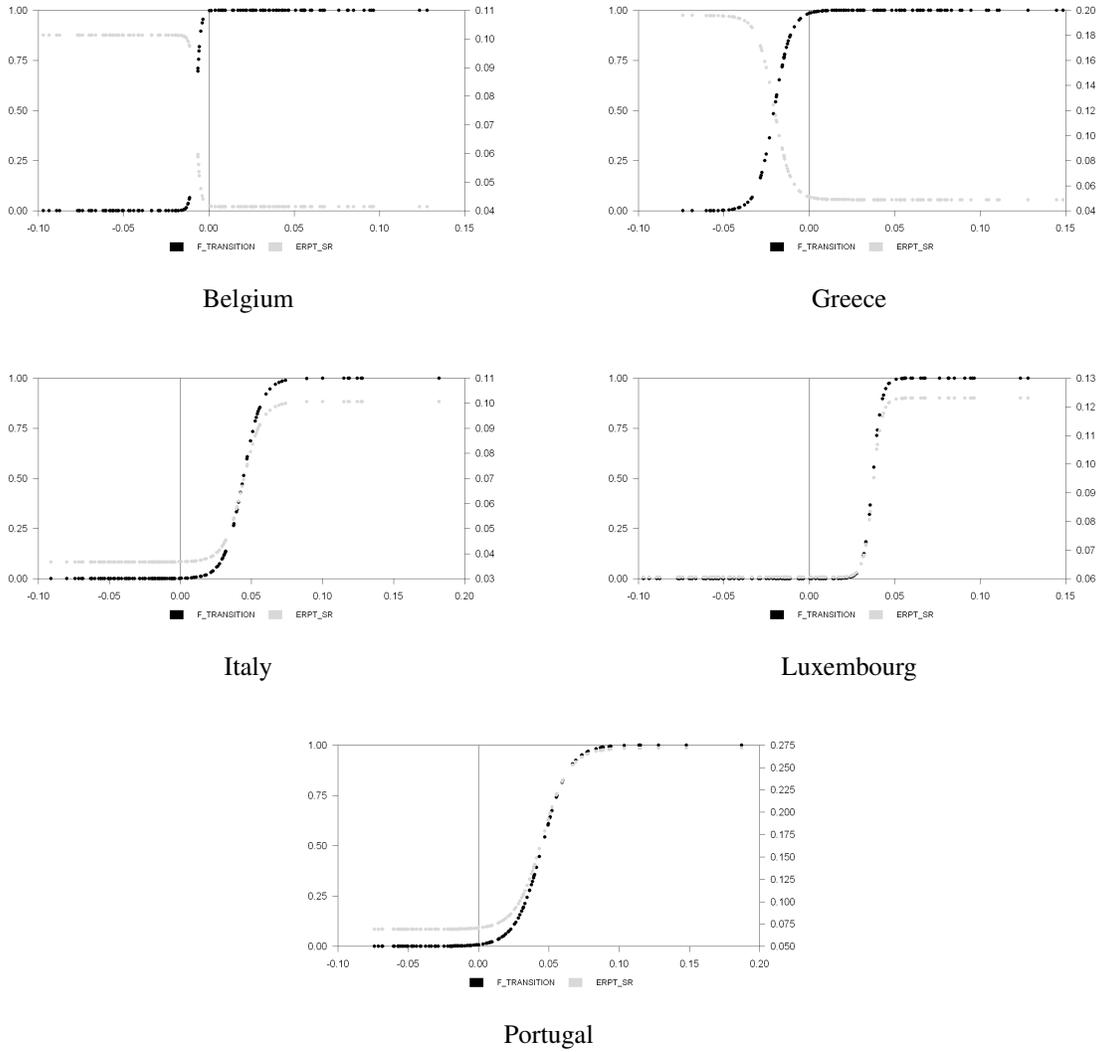
Note: Estimated transition function and short-run ERPT as a function of past inflation rates. Results are from LSTR with $s_t = \pi_{t-i}$.

Figure D.2: Logistic functions and long-run ERPT as a function of past inflation rates



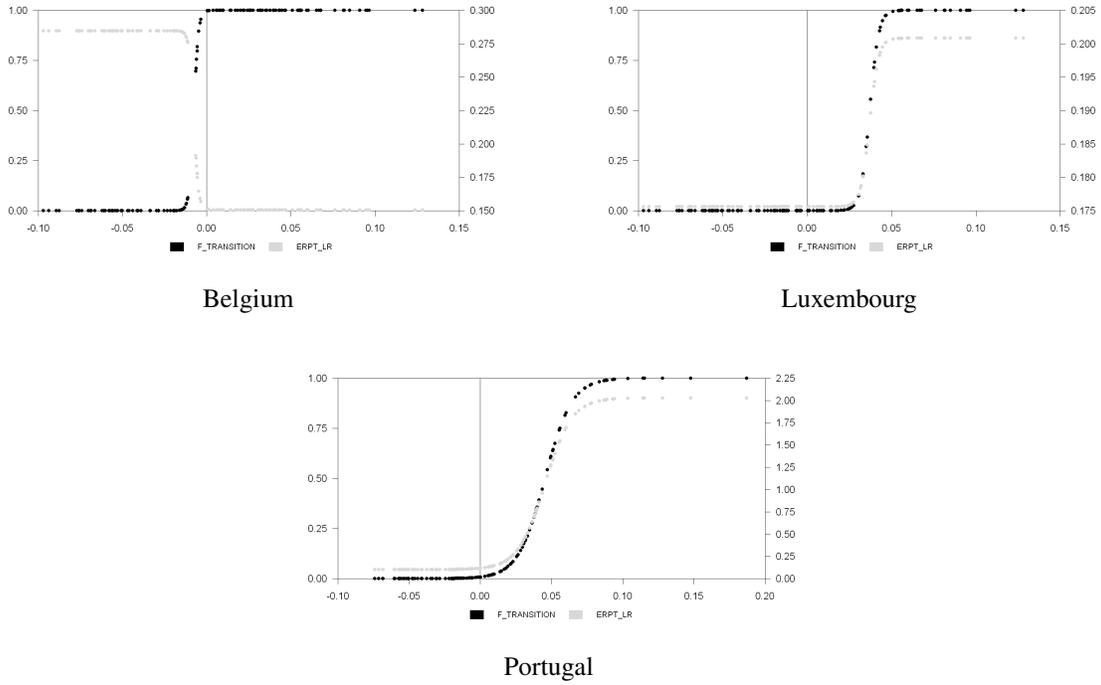
Note: Estimated transition function and long-run ERPT as a function of past inflation rates. Results are from LSTR with $s_t = \pi_{t-j}$.

Figure D.3: Logistic transition functions and short-run ERPT as a function of past depreciations



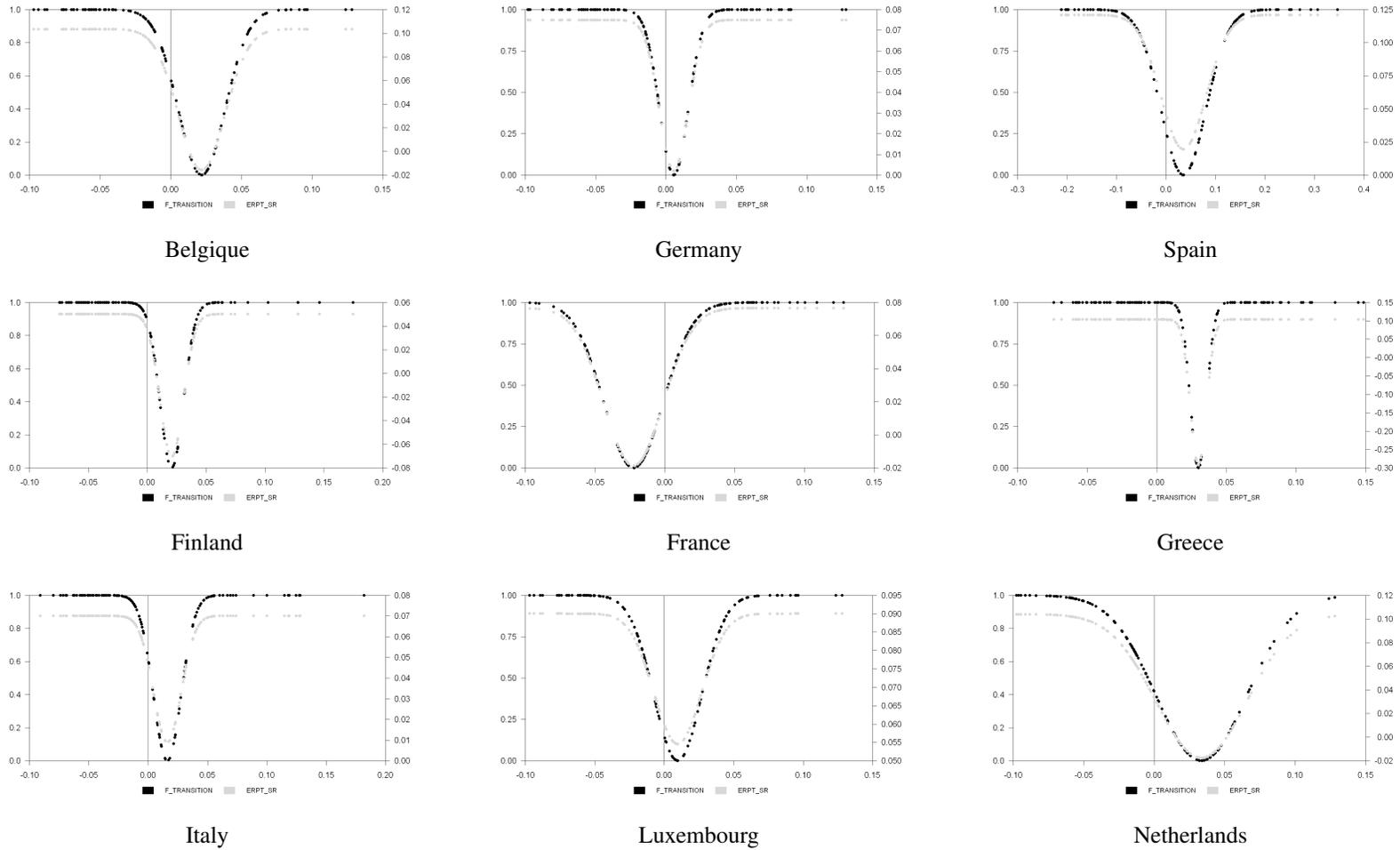
Note: Estimated transition functions and short-run ERPT as function of past exchange rate depreciations. Results are from LSTR model with $s_t = \Delta e_{t-i}$.

Figure D.4: Logistic transition functions and long-run ERPT as a function of past depreciations



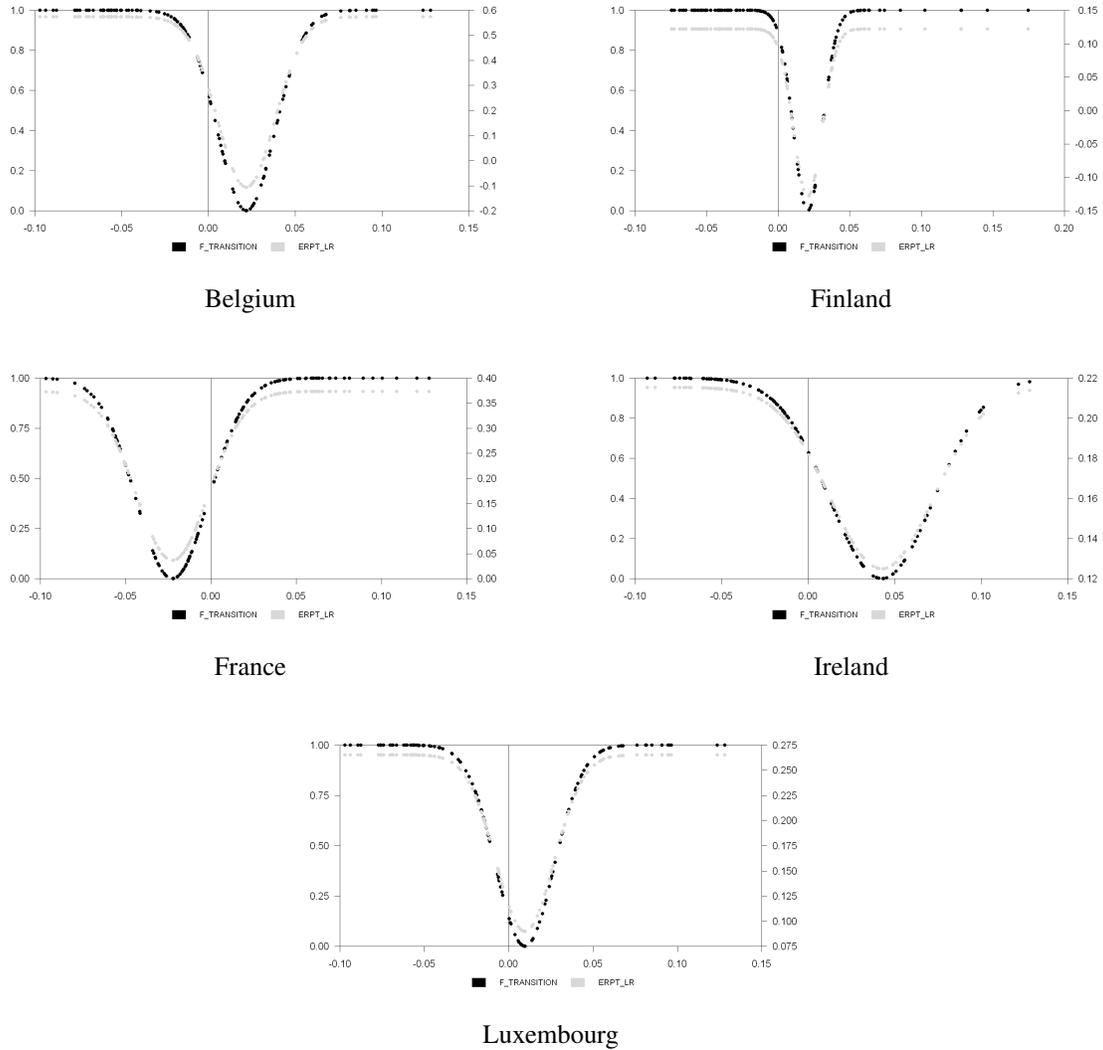
Note: Estimated transition functions and long-run ERPT as function of past exchange rate depreciations. Results are from LSTR model with $s_t = \Delta e_{t-i}$.

Figure D.5: Exponential functions and short-run ERPT as a function of past depreciations



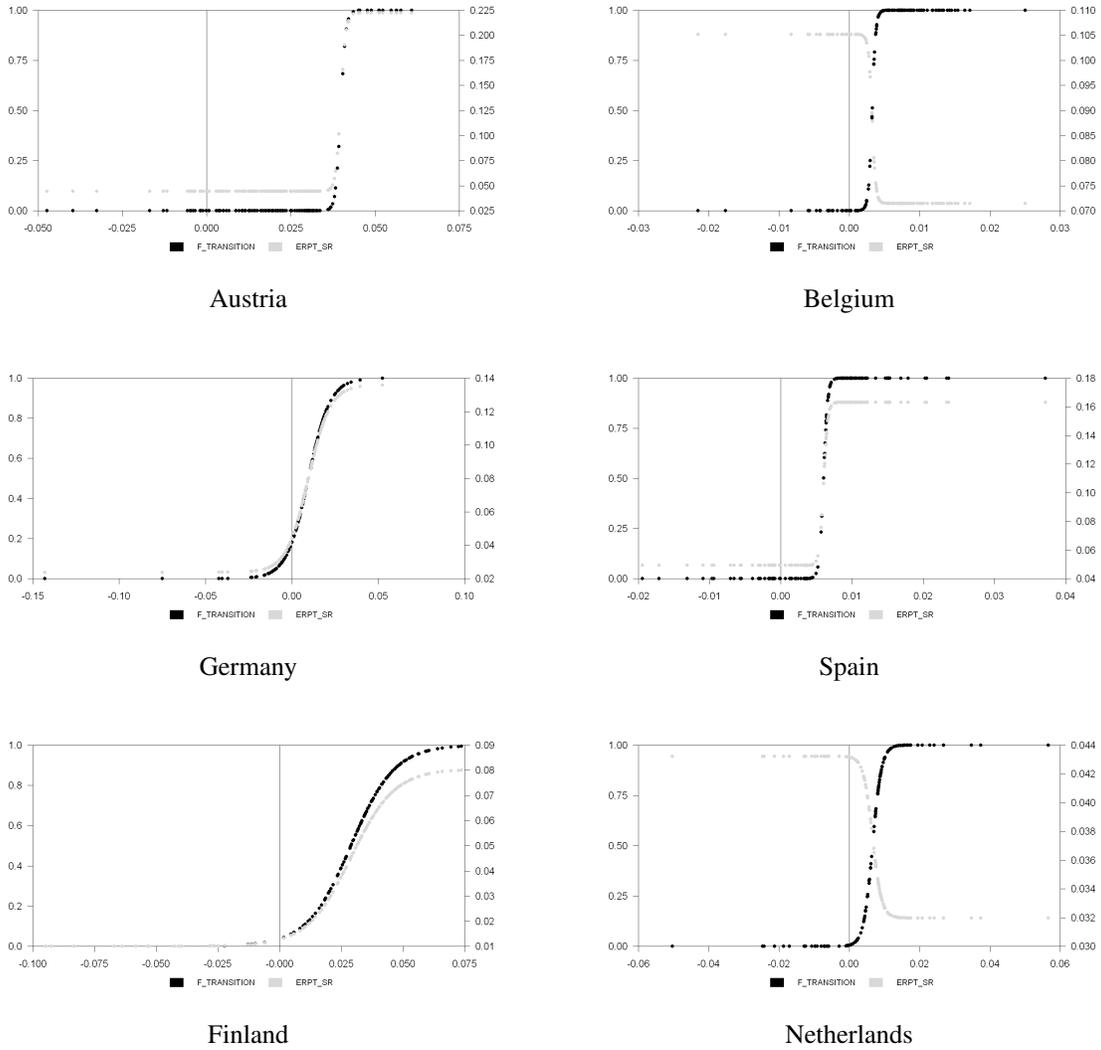
Note: Estimated exponential transition functions and short-run ERPT as a function of past exchange rate depreciations. Results are from ESTR specification with $s_t = \Delta e_{t-1}$.

Figure D.6: Exponential transition functions and long-run ERPT as a function past depreciation



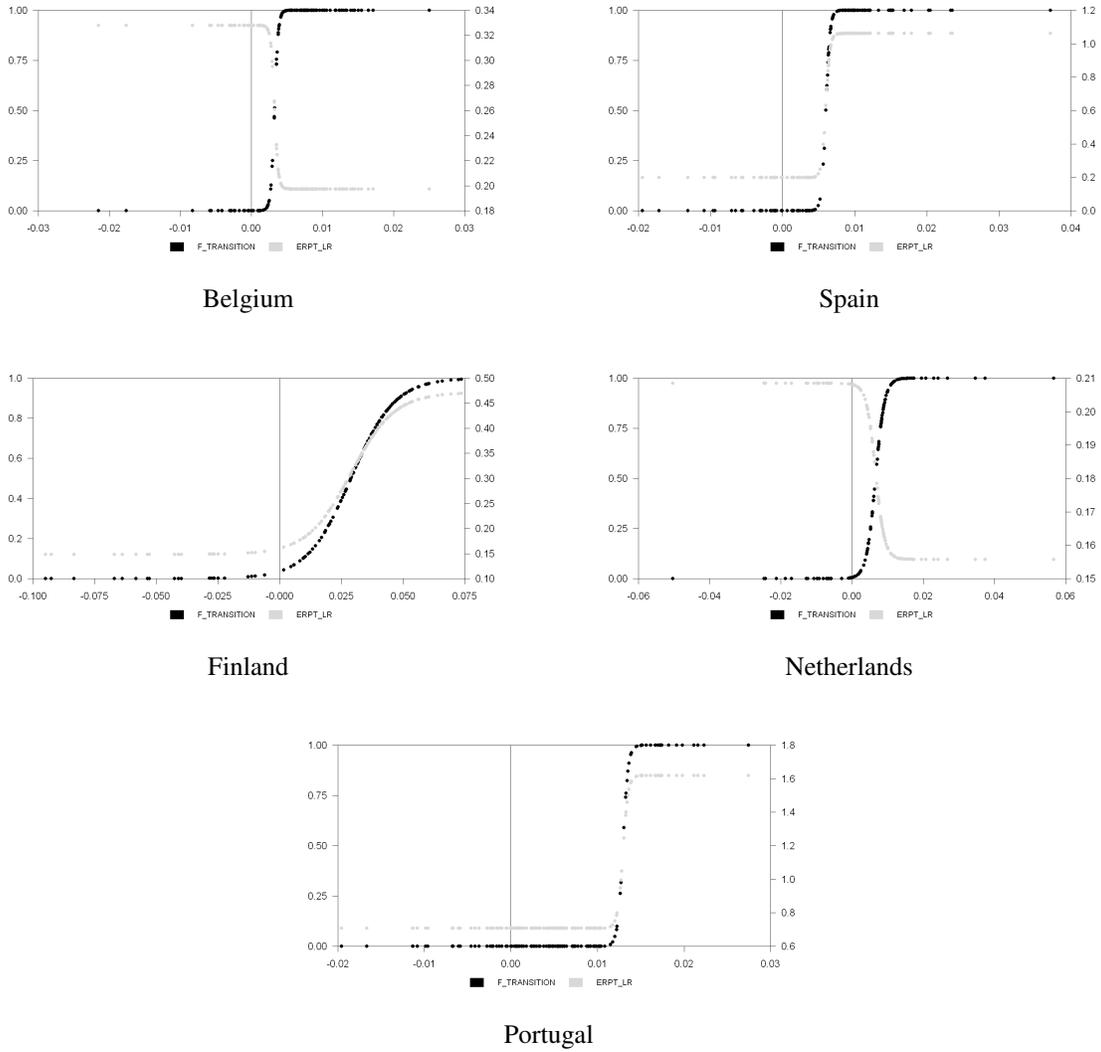
Note: Estimated exponential transition functions and long-run ERPT as a function past exchange rate depreciations. Results are from ESTR specification with $s_t = \Delta e_{t-i}$.

Figure D.7: Logistic functions and short-run ERPT as a function of past output growth

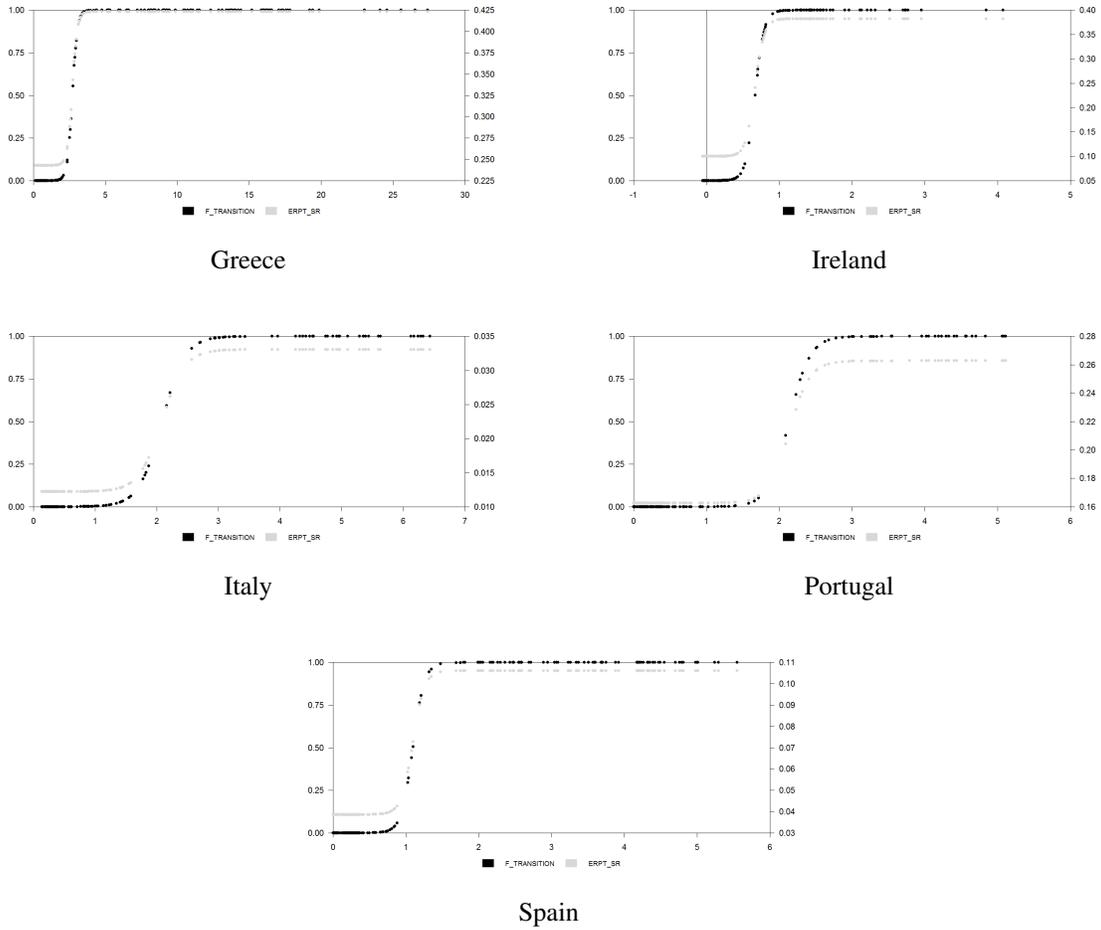


Note: Estimated logistic transition functions and short-run ERPT as a function of past output growth. Results are from LSTR model with $s_t = \Delta y_{t-i}$.

Figure D.8: Logistic functions and long-run ERPT as a function of past output growth

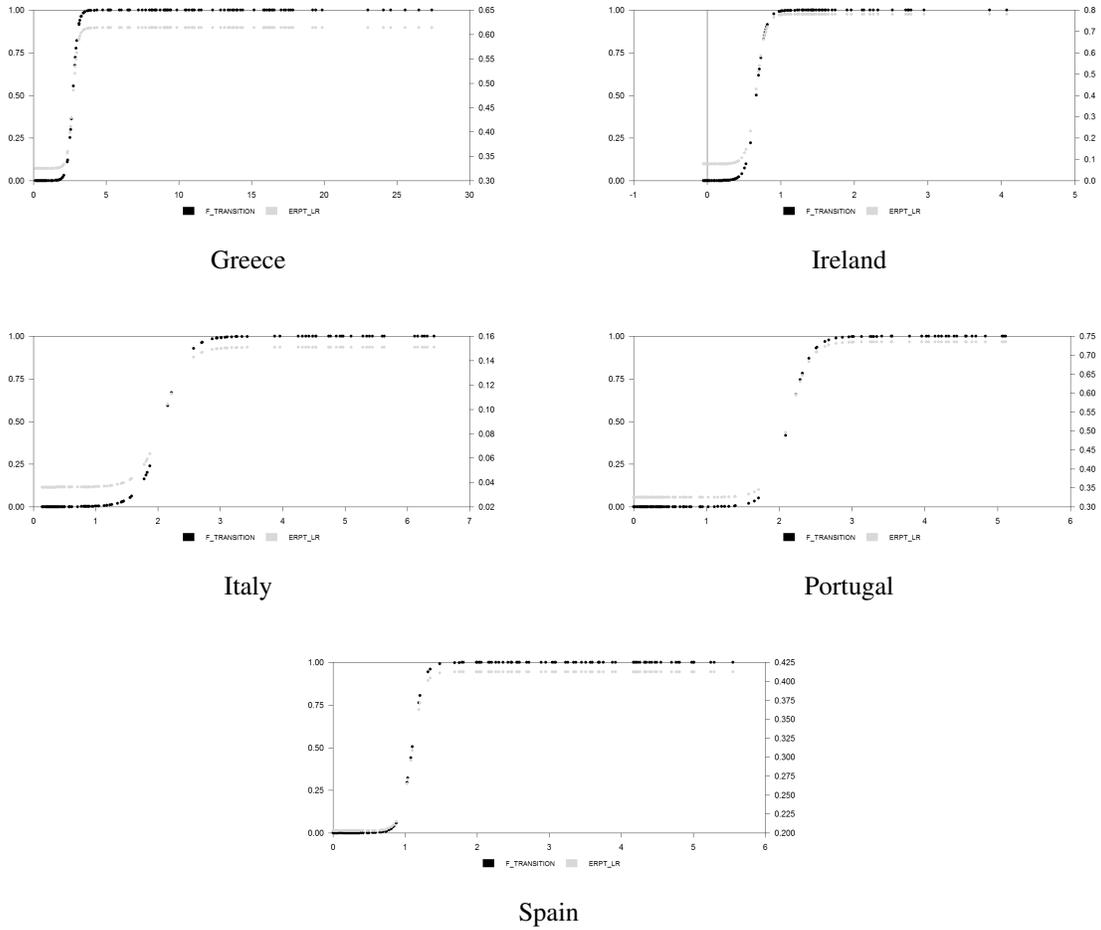


Note: Estimated transition functions and long-run ERPT as a function of past output growth. Results are from LSTR model with $s_t = \Delta y_{t-i}$.

Figure D.9: Logistic functions and short-run ERPT as a function of past yield spread

Note: Estimated transition functions and short-run ERPT as function of past bond yield spread. Results are from LSTR model with $s_t = bs_{t-j}$.

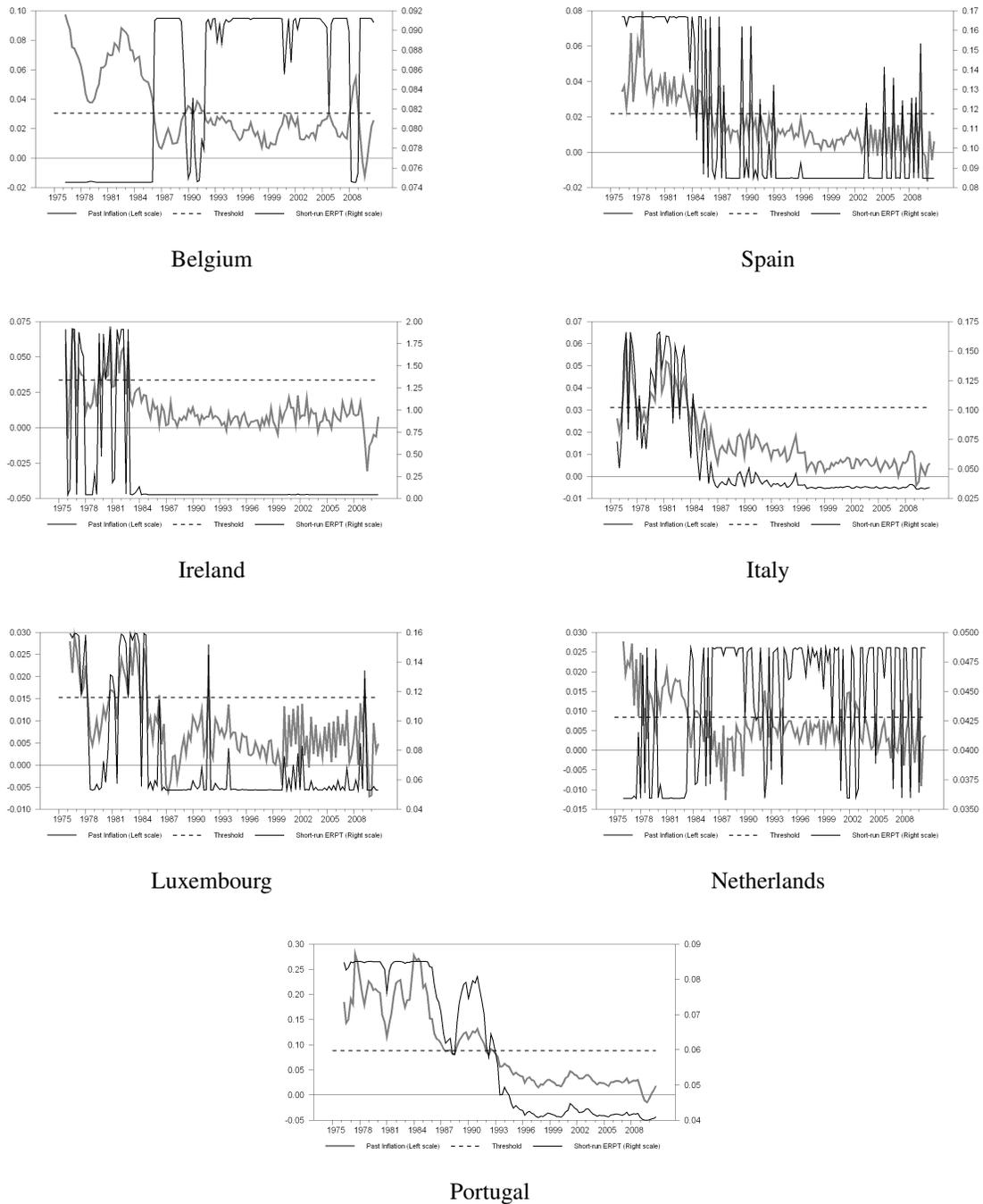
Figure D.10: Logistic functions and long-run ERPT as a function of past yield spread



Note: Estimated transition functions and long-run ERPT as function of past bond yield spread. Results are from LSTR model with $s_t = bs_{t-j}$.

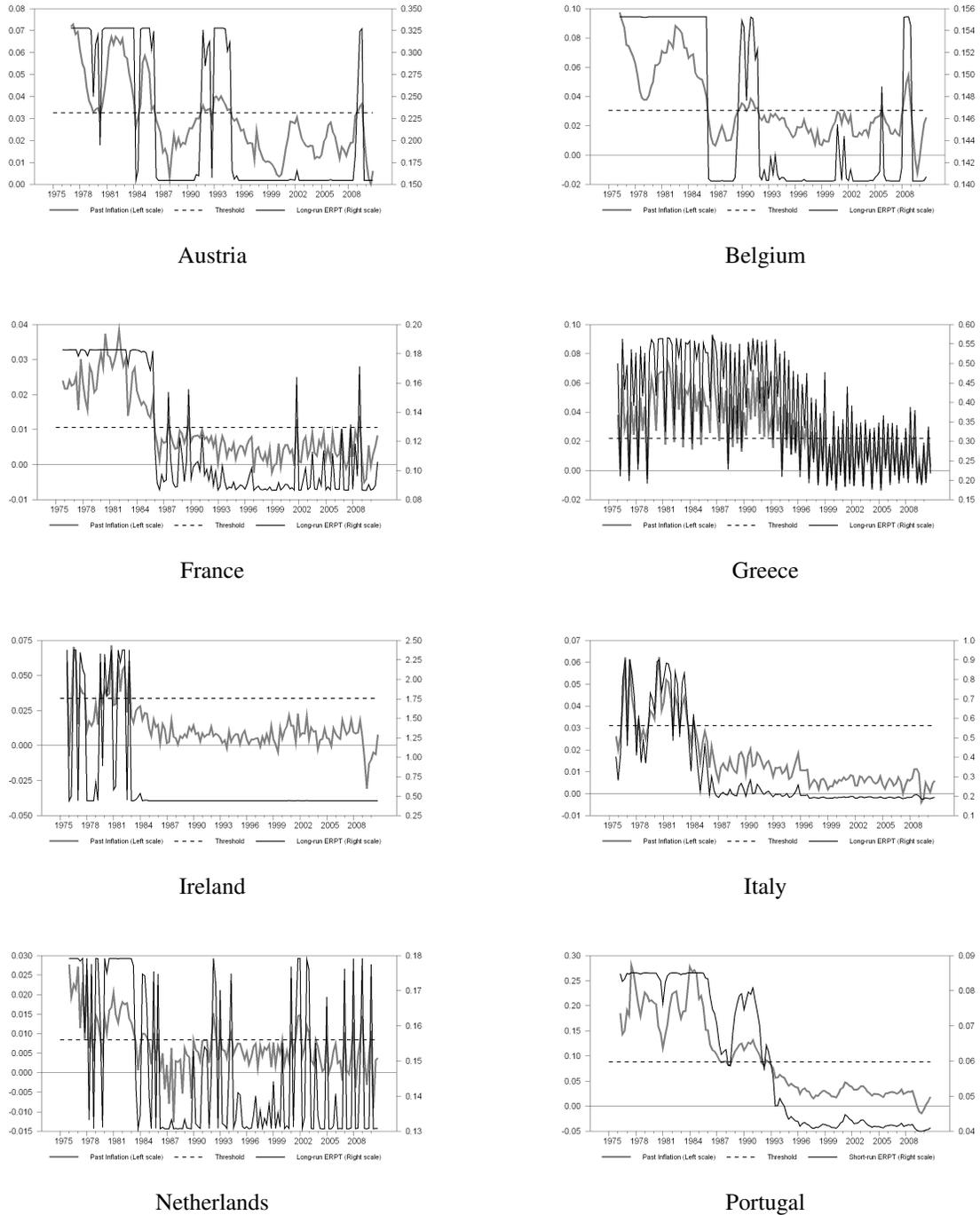
D.6. Plots of time-varying ERPT

Figure D.11: Time-varying short-run ERPT and past inflation



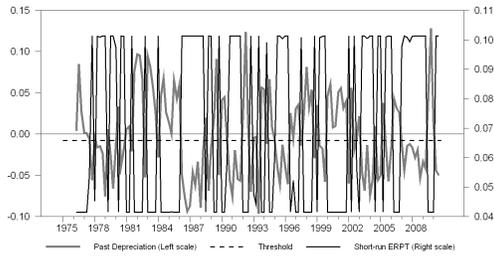
Note: Time-varying short-run ERPT and past inflation during 1975-2010. Results are from LSTR model with $s_t = \pi_{t-i}$.

Figure D.12: Time-varying long-run ERPT and past inflation

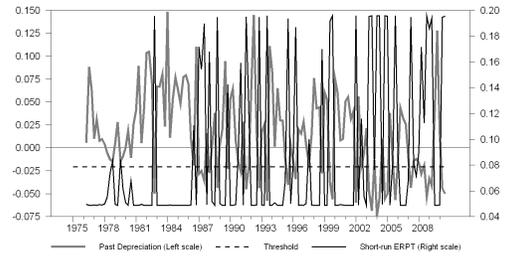


Note: Time-varying long-run ERPT and past inflation during 1975-2010. Results are from LSTR model with $s_t = \pi_{t-j}$.

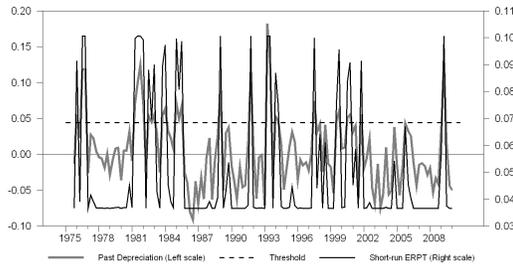
Figure D.13: Time-varying short-run ERPT and past depreciations (LSTR model)



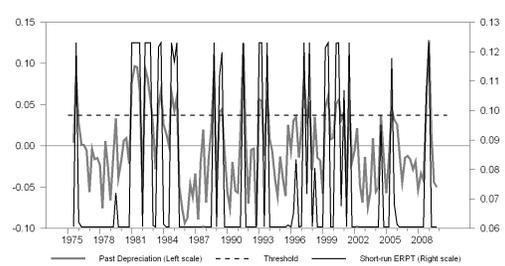
Belgium



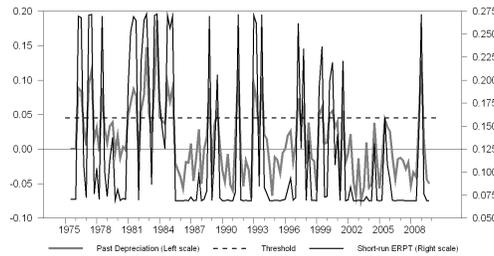
Greece



Italy



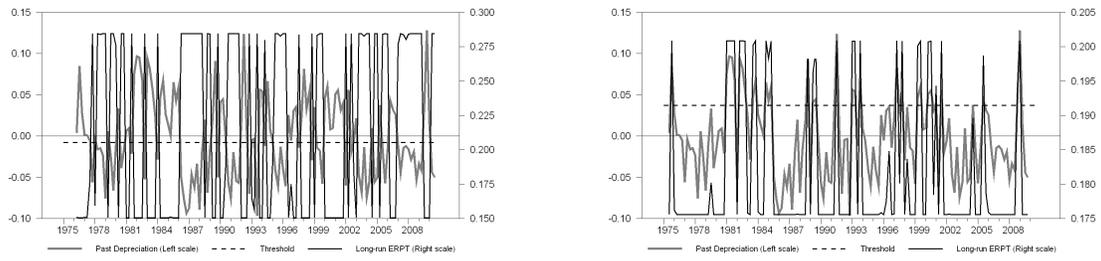
Luxembourg



Portugal

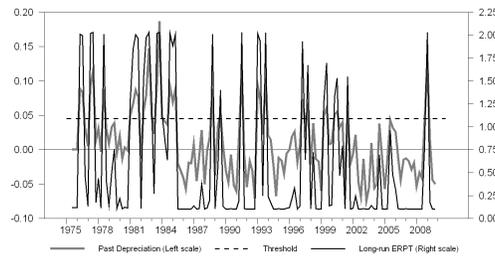
Note: Time-varying short-run ERPT and past currency depreciations during 1975-2010. Results are from LSTR model with $s_t = \Delta e_{t-j}$.

Figure D.14: Time-varying long-run ERPT and past depreciations (LSTR model)



Belgium

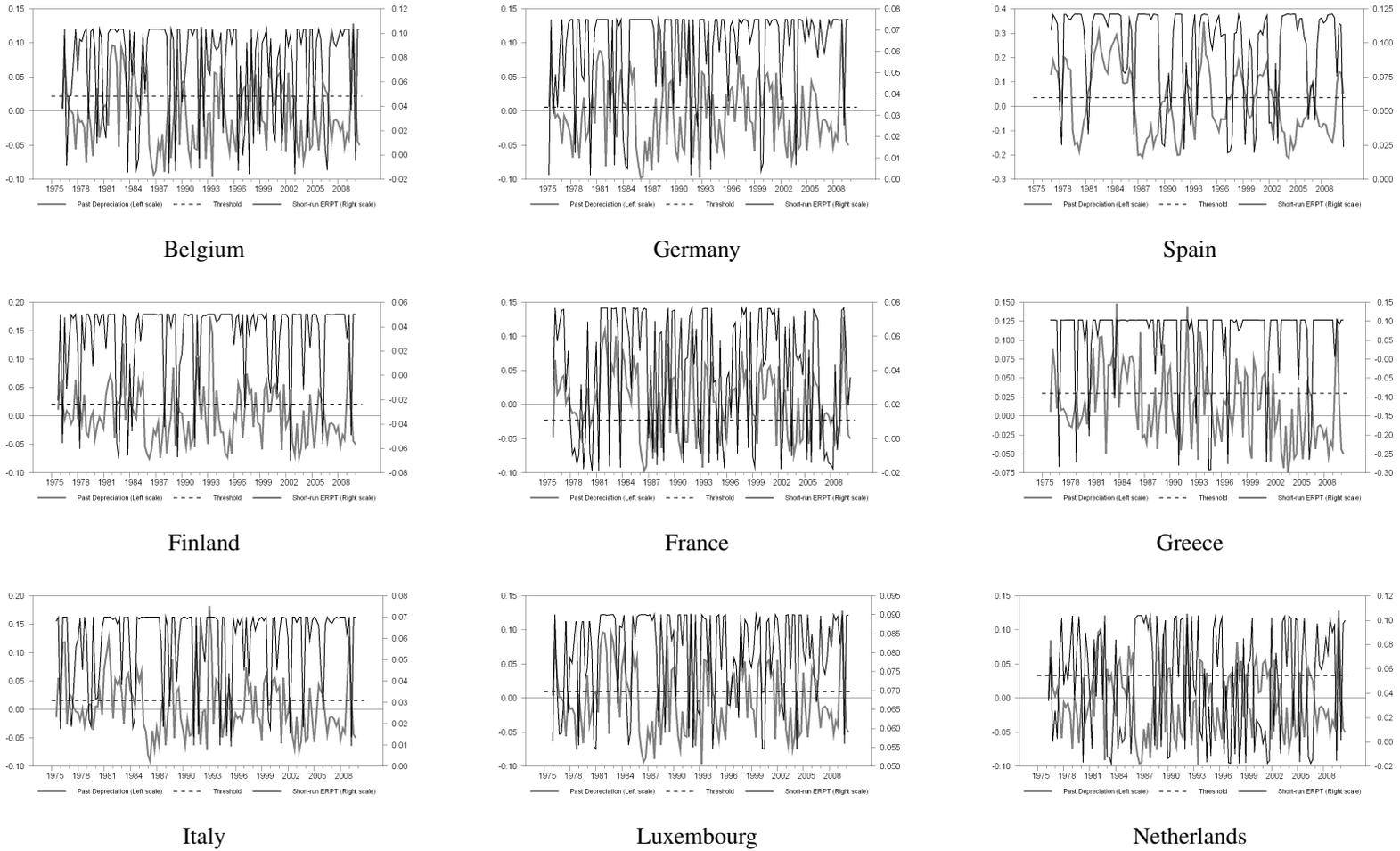
Luxembourg



Portugal

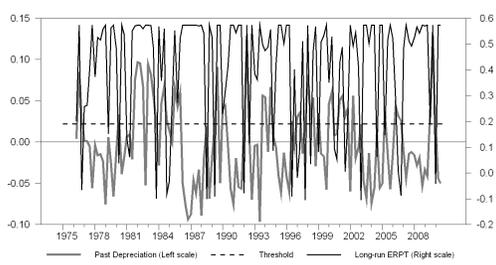
Note: Time-varying long-run ERPT and past currency depreciations during 1975-2010. Results are from LSTR model with $s_t = \Delta e_{t-j}$.

Figure D.15: Time-varying short-run ERPT and past depreciations (ESTR model)

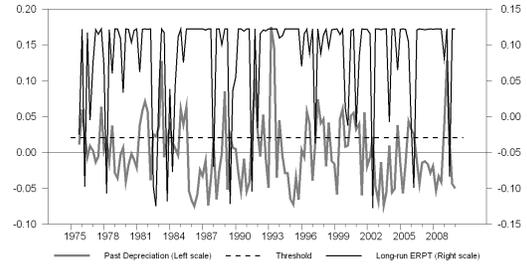


Note: Time-varying ERPT and past exchange rate depreciations during 1975-2010. Results are from ESTR specification with $s_t = \Delta e_{t-i}$.

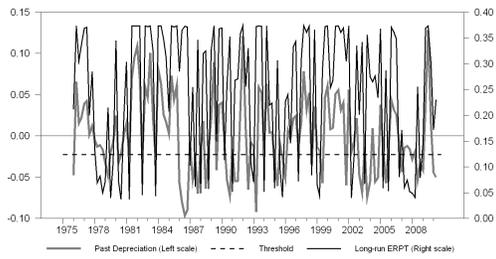
Figure D.16: Time-varying long-run ERPT and past depreciations (ESTR model)



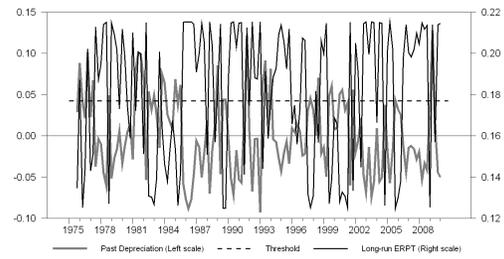
Belgium



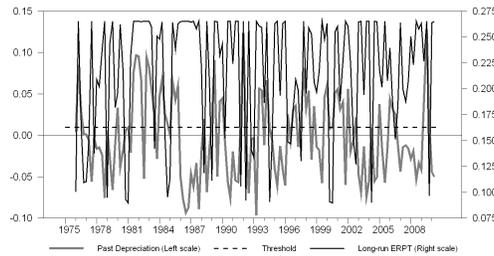
Finland



France



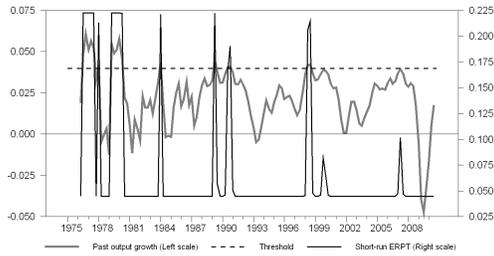
Ireland



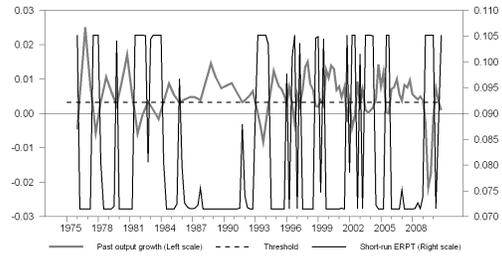
Luxembourg

Note: Time-varying long-run ERPT and past exchange rate depreciations. Results are from ESTR specification with $s_t = \Delta e_{t-i}$.

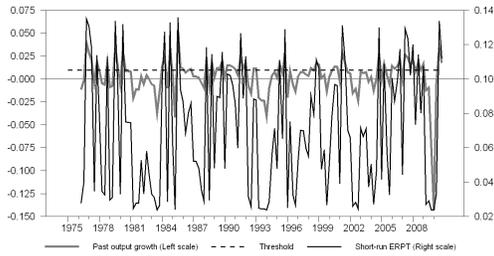
Figure D.17: Time-varying short-run ERPT and past output growth



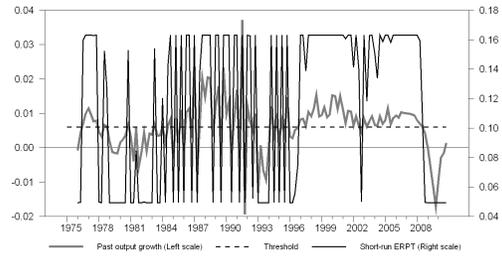
Austria



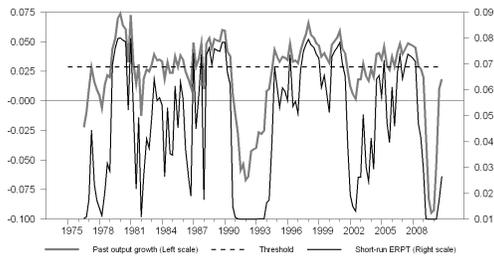
Belgium



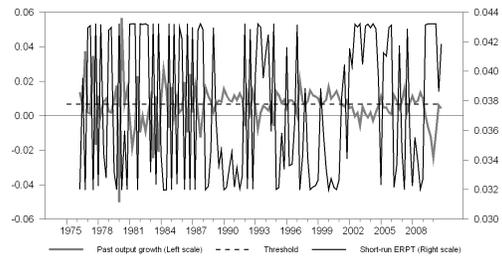
Germany



Spain



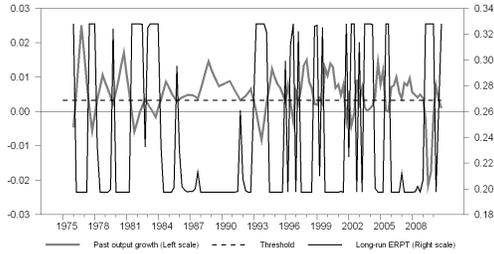
Finland



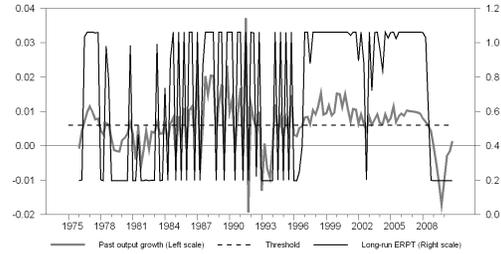
Netherlands

Note: Time-varying short-run ERPT and past output growth between 1975-2010. Results are from LSTR model with $s_t = \Delta y_{t-i}$.

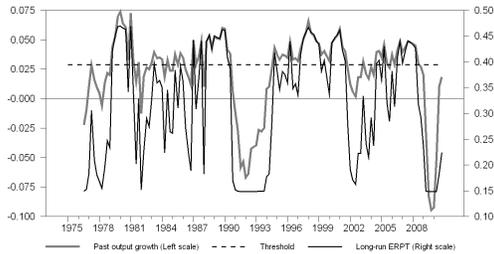
Figure D.18: Time-varying long-run ERPT and past output growth



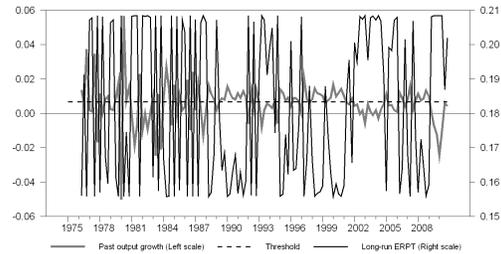
Belgium



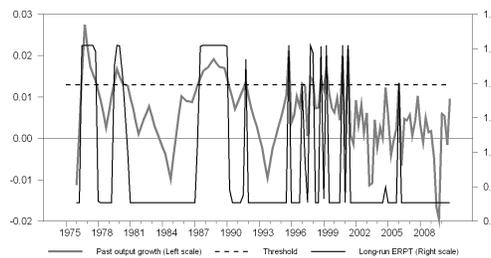
Spain



Finland



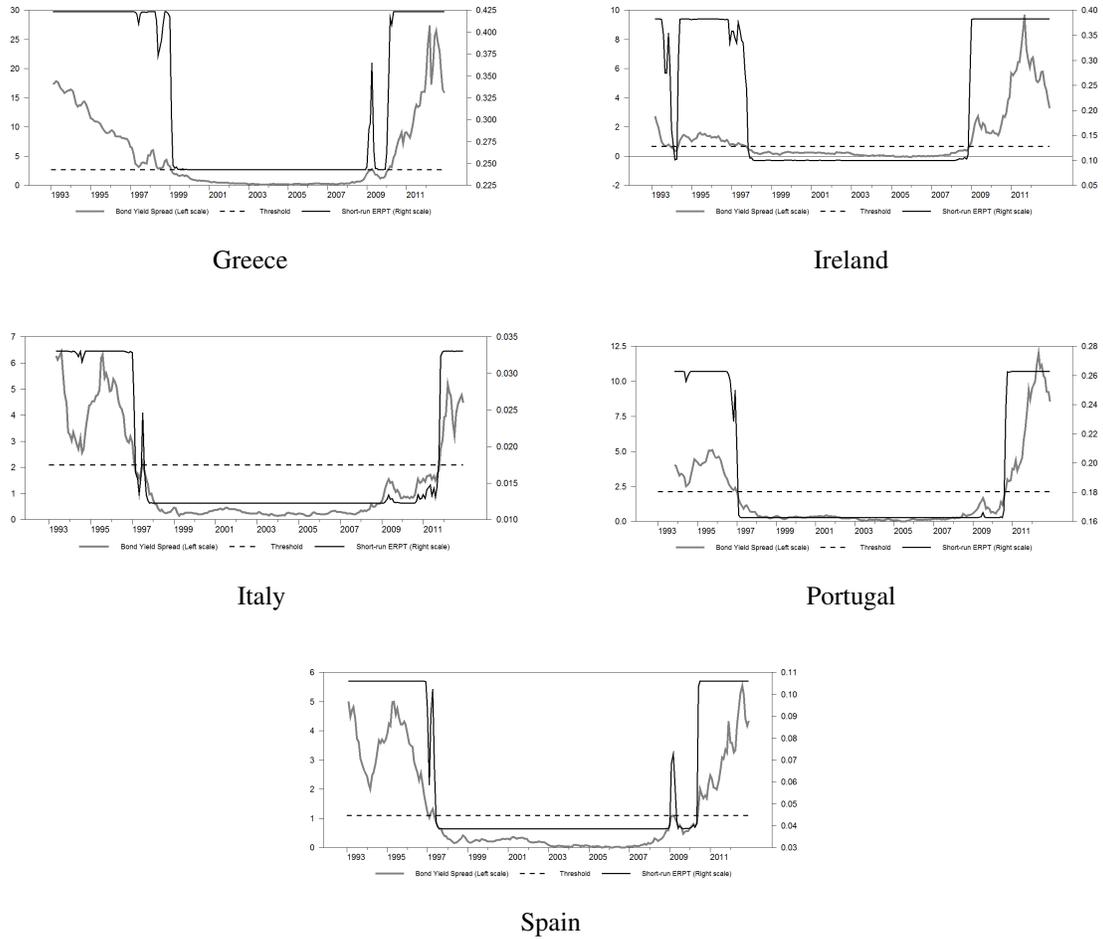
Netherlands



Portugal

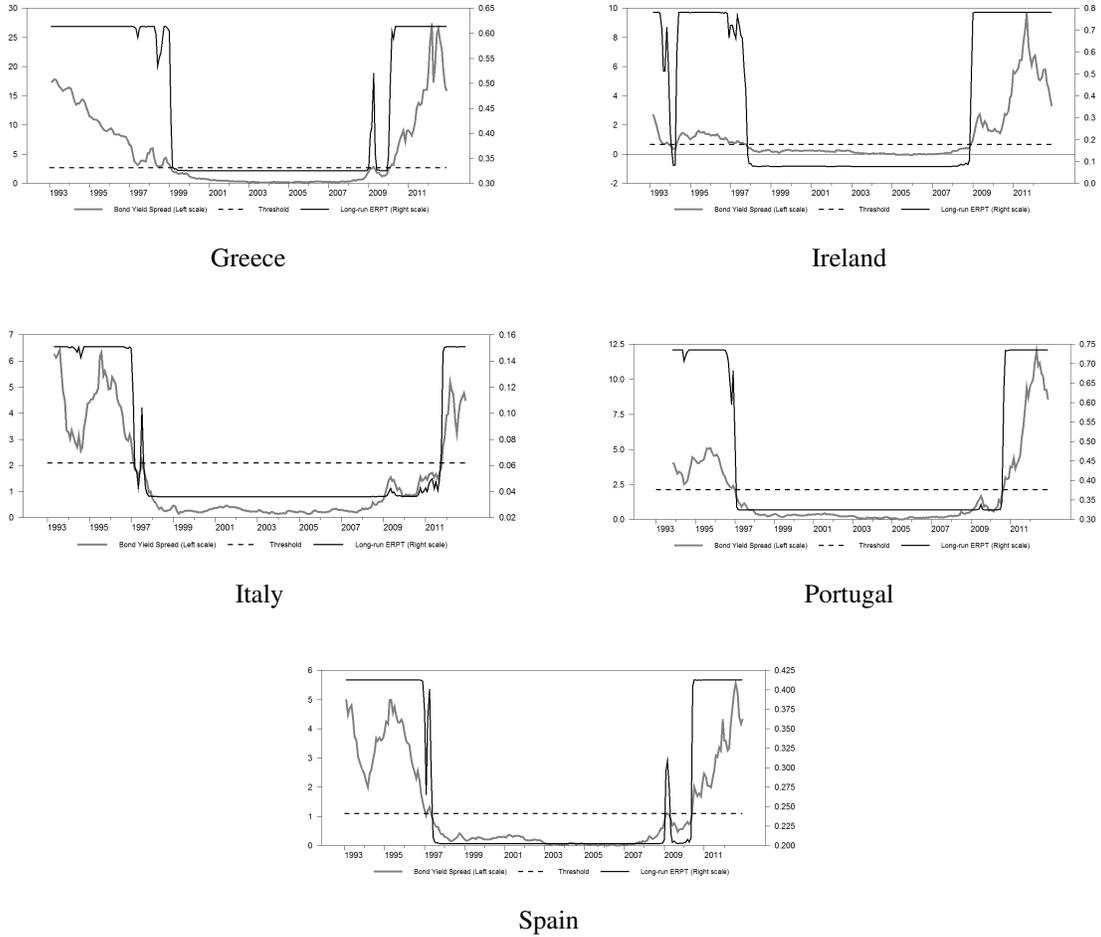
Note: Time-varying long-run ERPT and past output growth between 1975-2010. Results are from LSTR model with $s_t = \Delta y_{t-i}$.

Figure D.19: Time-varying short-run ERPT and past bond yield spread (LSTR model)



Note: Time-varying short-run ERPT and past bond yield spread during 1993-2012. Results are from LSTR model with $s_t = b y_{t-i}$.

Figure D.20: Time-varying long-run ERPT and past bond yield spread (LSTR model)



Note: Time-varying long-run ERPT and past bond yield spread during 1993-2012. Results are from LSTR model with $s_t = b y s_{t-i}$.

DISCUSSION AND CONCLUDING REMARKS

This thesis provides a **macroeconomic analysis** of the overall effect of exchange rate changes on the aggregate domestic prices over the last three decades, an issue of key importance for the conduct of monetary policy. The distinction between the pass-through of currency movements across import prices and consumer prices is crucial as they have different policy implications. As a matter of fact, it is desirable to have higher ERPT to import prices in order to use the exchange rate changes as an instrument for correcting the trade imbalances, while a lower pass-through to consumer prices is preferred as it avoids inflationary pressure to the domestic economy. Consequently, in *the first part* of our analysis we have focused on the ‘first-stage pass-through’, i.e. the responsiveness of import prices to exchange rate changes, and in *the second part* we have examined the transmission of the exchange rate movements via import prices to consumer prices. In this research, we have been motivated to study the euro area case as the introduction of the single currency since 1999 has been a more challenging task for the European monetary authorities to preserve price stability across the EA member states. The different exchange rate arrangements and the changing in macroeconomic conditions experienced by the Eurozone countries over time provide several implications that should be kept in consideration in the debate about the extent of pass-through. Also, we have employed a wide range of up-to-date econometric methods (dynamic panel data, panel cointegration, CVAR analysis, nonlinear STR models) in order to provide robust measures of the ERPT elasticities as well as to shed further light on the macro determinants of these pass-through coefficients.

The literature on ERPT has had a hard time to pin down with certainty the source of incomplete and declining rates of pass-through. As is well-known, this phenomenon has both macro- and microeconomic aspects, but the literature is not conclusive about

the most important factors, i.e. macro or micro factors. The distinction between macro- and microeconomic variables is very important since they point to substantially different implications in policy terms. If pass-through is a “macro” phenomenon which is directly associated with monetary policy, such as inflation or exchange rate volatility, this implies that a given decline in pass-through, may not necessarily be a permanent phenomenon because it may dissipate if monetary policy becomes more accommodative. In contrast, if ERPT is related to more structural factors, such as the industry composition of trade, economic policy is less capable to deal with. In fact, most of ERPT papers have argued the prevalence of micro factors in explaining the changing behaviour in ERPT. For instance, CAMPA and GOLDBERG (2005) argued that the macroeconomic variables - levels of inflation, money growth rates or country size - are weakly correlated with changes in pass-through, are not of first order importance in explaining pass-through evolution within the OECD countries over the past 25 years. By contrast, there is strong evidence that shift in the composition of imports toward goods whose prices are less sensitive to exchange rate movements has contributed to a fall in the pass-through in many countries in the 1990s. The same view was supported by GOLDBERG and TILLE (2008) who provide empirical evidence suggesting that the choice of invoicing currency is influenced more by the product composition of trade than by macroeconomic variables. Thus, the aim of our thesis is to contribute to the debate on the prevalence of microeconomic factors vs. macroeconomic factors. Moreover, it is worth noting that the majority of the earlier literature is micro-level studies, and in recent years it has been a fruitful area of research regarding the macroeconomic forces underlying pass-through. Hence, in our research we try to fill this gap by **highlighting the macroeconomic aspect of the ERPT** using relevant empirical methods.

Throughout the different empirical exercises, the main result of this thesis can be summarized as follows:

- In line with the bulk of literature, we provide a strong evidence of incomplete ERPT to the different domestic prices in the last two decades. The import-price pass-through, which is higher than in producer and consumer prices, is found to be incomplete even in the long-run. Also, the sensitivity of prices has declined markedly in recent years. For the EA countries, we find a steady decline in the degree of pass-through throughout the different exchange rate arrangements:

ERPT elasticity was close to unity during the *snake-in-the tunnel* period while it is about 0.50% since the formation of the monetary union.

- We find very weak evidence of decline around 1999, i.e. since the introduction of the single currency. Our results reveal that the pass-through estimates appear to trend down since the beginning of the 1990s. We notice that the observed decline was synchronous to the shift towards reduced inflation regime in our sample of EA countries.
- The substantial fluctuations in the exchange rate of the euro during the first three years of his existence has raised significantly the extent of pass-through. We pretend that this outcome would explain why the sensitivity of the import prices did not fall since 1999. Within VAR-system and dynamic panel data, we have pointed out that external factors, such as exchange rate shocks, had important inflationary impacts during 1999-2003, compared to the pre-EA period.
- Contrary to the previous empirical works, we don't find a wide heterogeneity in the degree of pass-through across the 12 EA countries during the last two decades. This may due to the process of monetary union which has entailed some convergence towards more stable macroeconomic conditions across member states. It is interesting to note that when we estimate our pass-through equation over 1979:2-1990:2, we point out more pronounced cross-differences in ERPT than recorded over 1990:3-2010:4. The divergent macroeconomic conditions across EA countries during the 1980s could be a potential explanation.
- There is an interesting distinction between *peripheral* and *core* EA economies in terms of pass-through. In general, the degree of ERPT to consumer prices appears to be most prevalent in Portugal and Greece, while the lowest coefficients were found in Germany, Finland and Austria. The historical path of inflation in each group of countries may explain this distinction. Countries with historically higher levels of inflation would experience higher rates of pass-through. Moreover, in the recent context of the European debt crisis, the sensitivity of consumer prices is found to be higher in the peripheral countries or in the so-called GIIPS group (Greece, Ireland, Italy, Portugal and Spain) which is mainly due to the general weakness of macroeconomic fundamentals.

- We find a substantial role for the macroeconomic variables as key determinants of the sensitivity of prices to currency fluctuations. Inflation regime, the exchange rate volatility and persistence, the magnitude of the exchange rate changes and the stability of the macroeconomic environment are the main factors influencing the degree of ERPT. While, there is no significant role for the degree of openness, measured as the ratio of imports to GDP. Within disaggregated data, it appears that the product composition of imports would determine the aggregate ERPT of an economy, and thus, cross-country differences in pass-through rates may be due to heterogeneous industry composition of trade across countries. However, unlike CAMPA and GOLDBERG (2005) hypothesis, we pretend that the trade composition of imports may explain some, but certainly not the lion's share of the decline in pass-through.
- Our study provides clear support for the presence of several sources of nonlinearities in the ERPT mechanism. Although there are different factors that may lead to nonlinear exchange rate transmission, this issue is routinely disregarded in most of the empirical implementations, and the few studies on this field has failed to find strong evidence of nonlinearity. In this thesis we have shown that the pass-through may be nonlinear with respect to number of macroeconomic variables, namely the inflation environment, the size of exchange rate changes and the macroeconomic instability.

Overall, we can say that exchange rate changes continue to lead to significant pressures on domestic prices in the context of the EA, although the general decline of ERPT. The wide swings of the euro during the first years of the EMU were a good illustration that the imported inflation remains a threat for the price stability, impacting differently inflation levels in the different member states. The main finding our thesis is the prominent role of macroeconomic forces in driving the extent of pass-through. A direct consequence of this result is that a changing behavior in ERPT is not necessarily a structural phenomenon and it may be solved via macroeconomic policies. The pass-through is also a macro phenomenon which is directly associated with monetary policy outcome, such as inflation environment. Thus, the shift to a more stable monetary policy conditions with credible and anti-inflationary regime within the EMU - as the main objective of the monetary policy of the ECB is to achieve medium-term price stability

for the EA aggregate - tend to reduce the degree to which the currency changes are transmitted to domestic prices. This is especially true for countries with historically poor macroeconomic policies, such as Greece, Ireland, Italy, Portugal and Spain. A better macroeconomic management and the adoption of a sounder set of policies within the monetary union may be an effective tool for reducing ERPT, and hence for decreasing the costs of maintaining inflation stability during turbulent exchange rate periods.

We conclude by indicating some limitations of the study and the possible directions for future research. In our exploration of the ERPT issue, we have followed an econometric approach with various methods, techniques and data specifications. Also, most of the frameworks described in this thesis are partial-equilibrium; exchange rates, input prices, costs were all taken as exogenous. It would be desirable eventually to integrate insights from the partial-equilibrium literature to general-equilibrium models that would inform monetary policy. In the recent years, there has been a burgeoning literature dealing with the ERPT in the context of new open-economy macroeconomics (NOEM) models. These latter are based mainly on work by OBSTFELD and ROGOFF (1995) which incorporates nominal rigidities and market imperfections into a dynamic stochastic general equilibrium (DSGE) models for open economies. The main advantage of these models is that, as in VAR models, avoid bias problem due to the endogenous variable, and thus take into account the fact that prices and exchange rate are determined simultaneously. Moreover, the DSGE framework allows analysis of the extent of pass-through conditional on specific shocks (e.g. monetary or productivity shocks.). Rather than assuming exchange rate changes are exogenous shocks that affect prices, the comovement between prices and the exchange rate may depends on the source of the shock. Consequently, the rate of ERPT may be different depending on the nature of shocks hitting the economy. As explained in the NOEM literature, the responsiveness of import prices to exchange rate movements are underestimated in the single-equation regressions, compared to the DSGE models, owing to an econometric bias related to the endogeneity of the exchange rate. We see the issues of the DSGE models as meriting further exploration in future research. It is an important task for empirical research is therefore to discriminate between the alternative models and econometric procedures in order to provide valid and reliable ERPT coefficients. Furthermore, the availability of data continues to be a challenge for researchers. The implication of this was that

some variables listed in the empirical model had to be proxied or constructed. For instance, the literature is undecided about which variable should be used as a proxy for the exporters' marginal costs. Some papers using broad consumer or wholesale price indexes and others unit labor costs (ULC) series derived implicitly from the nominal and real effective exchange rate as proxies for foreign firms costs (see e.g. CAMPA and GOLDBERG, 2005; BAILLIU and FUJII, 2004). As a matter of fact, such indexes may have become unsatisfactory proxies for costs of production, in part because of their relative insensitivity to movements in the prices of raw materials and energy or rather their higher volatility. Thus, the downside of such measures is that the model may not accurately capture the phenomenon under review, and can substantially affect the actual behaviour of pass-through. Looking for relevant measure of exporters' costs at the aggregate still required considerable effort and remains an important subject for future research. Another important subject for future research includes investigating the effect of exchange rate at a more disaggregated level using industry or sector-specific data. For instance, COUGHLIN and POLLARD (2004) argued that the mixed results regarding the direction of asymmetry in ERPT highlights the importance of analyzing pass-through at the industry level. If the direction of asymmetry varies across industries then aggregation may obscure asymmetry that is present at the industry level.

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La Transmission des Variations du Taux de Change aux Prix dans la Zone Euro

Résumé:

Cette thèse met en évidence l'aspect macroéconomique du degré de report du taux de change sur les prix dans la zone euro. Nous utilisons un large éventail de méthodes économétriques récentes afin de fournir des mesures robustes sur la transmission du taux change ainsi que sur ses déterminants macroéconomiques. Notre recherche révèle le rôle prépondérant des facteurs macroéconomiques dans le déclin récent du degré de report. Une conséquence directe de ce résultat est que la baisse du taux de transmission du change n'est pas nécessairement un phénomène structurel, et il peut être ainsi résolu par des politiques macroéconomiques conjoncturelles. Par exemple, l'adoption de régimes de politique monétaire plus crédibles avec l'engagement de maintenir une inflation faible joue un rôle important dans la réduction de la sensibilité des prix aux variations du change. Ceci est particulièrement valable pour les pays dont les politiques macroéconomiques sont historiquement laxistes. Ainsi, la poursuite de politiques économiques solide et lisible au sein de l'UEM peut être un outil efficace pour réduire le degré de report du taux de change.

Mots-clés: Degré de report du taux de change, prix des importations, prix à la consommation

The Pass-Through of Exchange Rate Changes to Prices in the Euro Area

Abstract:

This thesis highlights the macroeconomic aspect of the exchange rate pass-through to domestic prices in the euro area countries. We use a wide range of up-to-date econometric methods in order to provide robust measures of the rate of pass-through as well as to shed further light on its macro determinants. The main finding of our research is the prominent role of macroeconomic forces in driving the recent decline of the transmission of currency movements. A direct consequence of this result is that the lowering in the rate of pass-through is not necessarily a structural phenomenon and it may be solved via macroeconomic policies. For instance, the shift to a more stable monetary policy conditions with credible and anti-inflationary regime would reduce the sensibility of prices to exchange rate changes. This is especially true for countries with historically poor macroeconomic policies. Thus, a better macroeconomic management with a sounder set of policies within the EMU may be an effective tool for reducing the degree of pass-through.

Keywords: Exchange Rate Pass-Through, import prices, consumer prices