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PARIS SCHOOL OF ECONOMICS
UNIVERSITÉ PARIS I PANTHÉON SORBONNE

École doctorale Économie Panthéon Sorbonne
Laboratoire: UMR Paris-Jourdan Sciences Économiques

THÈSE

Pour l'obtention du titre de Docteur en Sciences Économiques

Présentée et soutenue publiquement le 2 juillet 2020 par

Nicolas JANNIN

ESSAIS EN ÉCONOMIE PUBLIQUE ET ÉCONOMIE GÉOGRAPHIQUE

Sous la direction d'Antoine BOZIO et de Stéphane GAUTHIER

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For the title of Doctor of Philosophy in Economics

Prepared and defended at the Paris School of Economics on July 2nd 2020 by

Nicolas JANNIN

ESSAYS IN PUBLIC FINANCE
AND ECONOMIC GEOGRAPHY

Under the supervision of Antoine BOZIO and Stéphane GAUTHIER

Jury:

<i>President</i>	Gabrielle FACK	Professor, PSE & Paris Dauphine University
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Table of Contents

- Remerciements v

- General Introduction 1

- 1 This Town Ain't Big Enough? Quantifying Public Good Spillovers 11**
 - 1.1 Introduction 12
 - 1.2 Institutional Background 16
 - 1.3 Theoretical Model 20
 - 1.3.1 Preferences 21
 - 1.3.2 Conditional Housing Demand 23
 - 1.3.3 Demand for Jurisdictions 23
 - 1.3.4 Housing Supply 24
 - 1.3.5 Labor Demand 25
 - 1.3.6 Public Good Supply 25
 - 1.3.7 Equilibrium 26
 - 1.3.8 Residential Amenities 27
 - 1.4 Data 29
 - 1.5 Reduced-Form Evidence 32
 - 1.5.1 DiD Framework 33
 - 1.5.2 Results 35
 - 1.5.3 Robustness Checks 46
 - 1.6 Structural Estimation 47
 - 1.6.1 Generalized Method of Moments 47
 - 1.6.2 Non-Parametric Evidence 48
 - 1.6.3 Moment Conditions 52

1.6.4	Estimation Results	53
1.7	Welfare Implications	58
1.8	Conclusion	61
2	Optimal Spatial Policies with Public Goods and Unobserved	
	Location Preferences	63
2.1	Introduction	64
2.2	Data	69
2.3	Stylized Facts on Public Good Agglomeration Economies	71
2.3.1	Raw Patterns	71
2.3.2	Descriptive Regressions	73
2.3.3	Preliminary Comments on Welfare	76
2.4	Economic Geography Model with Local Public Goods	77
2.4.1	Central Government	78
2.4.2	Demand for Cities	79
2.4.3	Demand for Private Goods	80
2.4.4	Supply and Ownership	80
2.4.5	Demand for Public Goods and Tax Competition	82
2.4.6	Equilibrium	83
2.5	Optimal Policies	84
2.5.1	Intuition in a Two-Region Example	84
2.5.2	Efficient Allocations	87
2.5.3	Optimal Transfers	90
2.5.4	An Efficiency Test	92
2.5.5	Efficiency of Observed Transfers	94
2.6	Equity and Density	97
2.6.1	Compensation and Responsibility	98
2.6.2	Revealed Social Preferences	98
2.7	Conclusion	102
3	The Deadweight Loss of Property Transaction Taxes	105
3.1	Introduction	106
3.2	Institutional Background and Data	109

3.2.1	French Administrative Geography	109
3.2.2	The French Stamp Duty	110
3.2.3	The 2014 Stamp Duty Reform	111
3.2.4	Data and Descriptive Evidence	113
3.3	Bunching	116
3.3.1	Bunching at the Time Notch	116
3.3.2	Bunching at the Border Notch	120
3.4	Extensive Responses	122
3.4.1	Treatment, Control and Spillovers	122
3.4.2	A Synthetic Control Approach	123
3.4.3	Average Effect on Number of Transactions	127
3.4.4	Average Effect on Prices and Quality	132
3.4.5	Statistical Inference	134
3.4.6	Public Spending	137
3.5	A Simple Search Model	138
3.5.1	Model	138
3.5.2	Welfare	141
3.6	Conclusion	144
	Bibliography	145
	List of Figures	155
	List of Tables	157
	A Appendix to “This Town Ain’t Big Enough? Quantifying Public Good Spillovers”	159
A.1	Jurisdiction Utility	159
A.2	Local Public Goods	159
A.2.1	Extension of the Model	159
A.2.2	Time Consistency and No Inherited Wealth	160
A.2.3	Taking the Extended Model to the Data	163
A.3	DiD Robustness Checks	165
A.4	Welfare	167

A.4.1	Before the Merger	167
A.4.2	After the merger	169
A.4.3	Welfare Comparison	171
A.4.4	Welfare Estimation	172
B	Appendix to “Optimal Spatial Policies with Public Goods and Unobserved Location Preferences”	173
B.1	Local Public Good Demand	173
B.2	A Two-Region Example	176
B.3	Planner’s Problem	180
B.4	Public Good Demand Calibration	182
B.5	Constant Expenditure Shares	185
B.6	Calibration of the Public Good Index	187
B.7	Skill Sorting	189
C	Appendix to “The Deadweight Loss of Property Transaction Taxes”	191
C.1	Comparative Statics	191
C.2	Proof of Proposition 1	192
C.3	Proof of Proposition 2	192
C.4	Proof of Proposition 3	193

General Introduction

This PhD thesis covers several topics in the fields of public finance and economic geography. The interplay of these two rich and dynamic fields is proving to be a fascinating area of research that allows to explore central questions pertaining to the optimal design of a nation's policies and institutions in their geographic dimension, both from a theoretical and an empirical perspective. In particular, the connection of these two research strands raises questions about the specific market or institutional failures related to the spatial aspects of public spending and taxes.

From a theory point of view, among the most pressing issues in normative public finance are the appropriate spatial scale at which taxing and spending political powers should be granted as well as the appropriate levels of local public goods and taxes to fund them — accounting for the welfare effects of induced behavioral distortions such as migrations or property market responses. Related, the optimal population size of cities and the desirable amount of spatial transfers are longstanding questions in urban economics. From an empirical perspective, the main challenge is to precisely quantify economic agents' behavioral responses to various forms of tax and transfer incentives, which link the micro-level individual actions at the root of various forms of market failures to the macro-level policy and institutional design that remedy them.

This thesis first revisits classic public finance topics — the provision of public goods and the welfare effects of taxes — and puts them in an economic geography context, an area where empirical evidence is scarce. Let's take the example of France. Following a series of decentralization laws starting in the early 1980's, French local authorities increasingly gained autonomy regarding local public services, being able to levy taxes and provide public goods to their constituents. The French local public sector, notably composed of around 35,000 local municipalities, together spent €229 billion in 2015, which represented a sizeable share of

GDP (around 10%), making local public spending and their welfare effects an area worthy of investigation. This spatial fragmentation and self-governance regarding public good provision raise specific questions about institutional efficiency and about the optimality of government policies interacting with this institutional context.

On the government policy front, it is only recently that urban economists have started applying long-developed theoretical insights (see [Kline and Moretti 2014](#) for a review) to actual data to study the welfare effects of large governmental place-based programs ([Kline and Moretti 2013](#); [Busso et al. 2013](#)) and spatial transfers more generally ([Albouy 2012](#); [Fajgelbaum and Gaubert 2018](#)). In the area of fiscal federalism and institutional design on the other hand, theory abounds (e.g., [Besley and Coate 2003](#)) but evidence that connects it to the data is scarce.

I first focus on an institutional question: the optimality of observed local jurisdictions' boundaries in the presence of public good spatial spillovers. Since the seminal works of [Tiebout \(1956\)](#) and [Oates \(1972\)](#), academics have investigated the optimal balance of power between local and central governments and the appropriate geographic level at which public goods should to be provided. Too much centralization causes decision-makers to lose touch with local needs, while too much decentralization creates geographic externalities as the benefits of public goods might spread beyond the boundaries of jurisdictions providing them. In the latter case, public goods may be under-provided from a collective welfare perspective. However, the literature lacks a framework for empirically evaluating these different effects, and this thesis proposes such one framework for evaluating cross-boundary benefits specifically.

I then study public good migration, taxation externalities and jurisdictions' optimal population sizes. Holding constant boundaries and institutions, additional sources of inefficiency related to the provision of public goods may require central policy intervention. Since the work of [Flatters et al. \(1974\)](#), several authors have studied the issue of scale economies in public good consumption and of migration and other local taxation externalities. I revisit these questions and examine how a government can make spatial transfers that achieve an efficient and equitable population distribution. Taking seriously the existence of location preferences sheds a different light on these issues by asking a simple question: when people are attached to places, for good or bad reasons, how much does it change how we think about spatial policies, efficiency and equity? The answer seems to be "quite a lot" and may call into question large infrastructure projects and urban development plans but also direct spatial transfers designed

at making dense and attractive places denser and even more attractive, at the expense of more rural areas for possibly no aggregate gains.

Finally, I investigate the functioning of local property markets and local taxes and go back to exploring a classic public finance question: what are the behavioral responses to taxation, here transaction taxes? Behavioral responses to taxation are pivotal to the funding of local public goods, as they limit the scope of what government can achieve with local taxation. Evidence on property markets to taxation is building up but is relatively recent, and this thesis makes a hopefully credible contribution to this new strand of research. Although lump-sum taxation should be preferred to distortive taxation to fund local public goods, property transaction taxes may have other motives as they may also be used to tackle other market failures specific to frictional property markets.

A Systematic Use of Exhaustive Data

The last two or three decades have seen a boom in the use of administrative data — in lieu of survey data — for empirical analysis, which greatly contributed to the “credibility revolution” in applied work together with better identification strategies (see [Angrist and Pischke 2010](#); [Kleven 2018a](#); [Currie et al. 2020](#)). This revolution has been very much apparent in labor and public economics with the growing availability of exhaustive individual-level microdata, but also in urban economics, with more and more databases with fine-grained spatial information being made available to researchers, including in France.

In an attempt to shed some light on the above-mentioned questions, this thesis takes advantage of rich administrative databases of various kinds — some publicly available, some regularly made available to researchers under strict conditions and some I obtained first-time access to — managed by the French administration that carries out a precious and indispensable work of data collection and cleaning. Most of these databases are used in more than one chapter across this thesis. They belong to five categories: property and population stock data, local public finance data, property transaction and housing prices data, local income and socio-demographic data, and administrative geography data.

Property and population stock data The FILOCOM (*Fichier des Logements à la Commune*) database is one of the cornerstones of this thesis and is used across the three chapters. It is an exhaustive dataset on household housing stock which provides information on each

non-commercial property in France every two years between 1994 and 2014. FILOCOM is managed by the French Ministry for the Environment which uses it for in-house studies. It is constructed by the Ministry of Finance’s Directorate General for Public Finances (*Direction Générale des Finances Publiques* or DGFIP) that merges exhaustive registers used for the calculation and management of local residence and property taxes, together with the national income tax register. FILOCOM contains information on each residential dwelling’s precise location, characteristics (surface, number of rooms, etc.) and vacancy or occupation status. Importantly, it also features the dwelling’s number of inhabitants. FILOCOM is hence a crucial — and the only — source for comprehensive high-frequency population and property stock information at the most local level in France.

Local public finance data Local tax and spending data are pivotal to this thesis, in particular to chapters 1 and 2. I first use a comprehensive database called *Balance Comptable des Communes*, obtained from the Institut des Politiques Publiques and the Ministry of Finance, covering the financial accounts of all 35,000 French municipalities and all 2,000 municipal federations for every year over the period 2002–2016. These rich datasets provide detailed information on jurisdictions’ public assets and liabilities, various spending items, local tax revenues, user and business fees as well as various subsidies and transfers received from higher jurisdictions.

I also exploit detailed local tax data (*Recensement des Éléments d’Imposition à la Fiscalité Directe Locale* — REI) for all 35,000 French municipalities and all 2,000 municipal federations for every year from 2002 to 2016, also obtained from the Institut des Politiques Publiques and the Ministry of Finance. REI contains detailed data on all tax bases, rates and revenues at the municipal and municipal federation levels for local household property and residence taxes as well as local property and non-property business taxes.

Property transaction and housing prices data This thesis uses several datasets related to property transactions, including the widely used databases collected and managed by the notary offices (BIEN database for the Parisian region and PERVAL database for the rest of France). BIEN and PERVAL contain transaction-level information including location, housing characteristics and transaction price. Geographic coverage is imperfect and notably suffers from under-reporting in the most rural areas in PERVAL. It is nonetheless a high-quality data source, in particular for housing prices, and is available every two years since 2000 from the

Ministry for the Environment.

In addition, I gained first-time access to *Demandes de Valeurs Foncières* (DV3F), a property transaction dataset obtained from the Ministry for the Environment for the period 2010–2016. DV3F is produced using other official databases including the Ministry of Finance’s transaction tax database and property register. It provides an exhaustive coverage of all transactions on the primary and secondary French property markets, that is all land and built property transactions involving natural persons, businesses or public entities. DV3F contains detailed information on prices, properties’ size and many other characteristics (location, number of living rooms, bedrooms, garages, etc.). It covers the whole French territory except for the 3 border counties — *Moselle*, *Bas-Rhin* and *Haut-Rhin* — that were lost to Germany during the 1870 Franco-Prussian war and reintegrated in 1919 after WWI.

Local income and socio-demographic data Across this thesis’ three chapters, I use a comprehensive dataset called IRCOM (*Impôt sur le Revenu par Commune*) obtained from the Ministry of Finance giving labor and social security total income at the municipal level from 2002 to 2016 for all 35,000 French municipalities. I complement it with publicly available INSEE census data to get further municipal-level information on socio-demographic characteristics for years 1990, 1999, 2008 and 2013.

Administrative geography data Finally, I use various publicly provided INSEE datasets on local jurisdictions’ geography (municipal federations’ membership history, jurisdictions’ creations, mergers and separations, geographic position within urban areas, neighbor jurisdictions, etc.). I also use detailed data on natural and urbanized land area at the municipal level for each year between 2008 and 2014 from the Ministry for the Environment.

Methods: Connecting Theory to the Data

Theory is most useful when it can be taken to the data in a transparent fashion. Throughout this thesis, I apply economic theory and modeling to try to answer policy and institutional questions and attempt to give flesh to these theoretical insights by making an extensive use of available databases and recent developments in econometrics.

Chapter 1 follows the footsteps of recent research in urban economics (e.g., [Ahlfeldt et al. 2015](#); [Diamond 2016](#)) and develops a structural framework, that is, a fully-fledged model

that explicitly features so-called deep or structural parameters capturing the fundamental architecture of preferences, technology and the environment that agents evolve in. This kind of model is easily amenable to empirical analysis in a transparent fashion, and estimation can be carried out using policy changes or external shocks of any size. However, the identification requirements are typically demanding — many big datasets and usually more than one source of exogenous variation — since these models typically rest upon a large set of parameters and variables. Once estimated, a structural model like this can be used to make welfare predictions for reforms of any type and size.

Away from fully structural approaches, chapters 2 and 3 use the public economics’ optimal taxation toolbox and characterize the optimal policies tackling specific market failures by the way of “sufficient statistics”, a term coined by [Chetty \(2009\)](#) whose use has exploded in the last decade. The sufficient statistics approach easily applies to setups beyond income redistribution, such as Pigovian taxes correcting non-governmental externalities. Its elegance lies in its simplicity: the welfare effects of marginal policy changes can be expressed using only a reduced set of observables, high-level parameters and elasticities that are robust to the model’s primitives — which can then rest upon quite general functional forms. However, it has shortcomings: the method can only be used to evaluate past (and small) reforms and complete characterisation of optimal policies or welfare diagnoses of observed situations may be beyond its reach without resorting to more structural approaches (see [Kleven 2018b](#) for a recent discussion on sufficient statistics).

This thesis’ attempts at connecting these models to the data build on the many additions to the empiricist’s toolbox over the last two decades, which also contributed to the identification revolution in applied econometrics together with better data. Designed-based identification strategies that rely on quasi-experiments such as difference-in-differences (DiD), event studies, bunching analysis and to a lesser extent more recent synthetic control methods, together with convincing graphical evidence — the “graphical revolution” — have been increasingly used since the early 1990s and are now pervasive in empirical work ([Currie et al. 2020](#)). At the same time, some economists have criticized these recent developments, pointing to the fact that applied work is allegedly increasingly focused on narrow if not second-order problems (see [Angrist and Pischke 2010](#) for a response to these criticisms).

In any event, chapter 1 is an example of how design-based empirical analysis can try to answer “big” questions pertaining to the optimal design of institutions. It provides a series of

graphical DiD evidence on the effects of local shocks on many local outcomes— together with further evidence on the decomposition of these shocks and their effects, providing additional variation— and uses these shocks and additional evidence as a basis for a General Method of Moments (GMM) estimation of the structural model. Importantly, the contact between theory and econometric analysis is never broken, and reduced-form as well as GMM specifications directly flow from the model. Chapter 3 uses bunching analysis and a synthetic control strategy. It gives bunching and DiD graphical evidence of behavioral responses to small tax changes across France’s mainland local jurisdictions. DiD estimates are then plugged into the sufficient statistics formula to evaluate the welfare effect of these past tax changes. While chapter 2 does not feature design-based identification, it takes sufficient statistics formulas (with some additional structural ingredients) to the data to characterize the desirability of marginal reforms in observed situations.

Outline and Summary

Chapter 1 was written with Aurélie Sotura and is entitled “This Town Ain’t Big Enough? Quantifying Local Public Goods Spillovers”. It is the first known attempt at structurally estimating local public good spillovers. In this chapter, we develop a simple and flexible framework that allows to test for the presence of public good spillovers between fiscally autonomous jurisdictions— i.e, how much residents of a given jurisdiction benefit from public goods provided by close-by jurisdictions— and to investigate potential welfare gains from increased centralization of public good provision. We build a quantitative spatial equilibrium model of cities with mobile households, imperfectly elastic housing supply and endogenous local public goods causing spillovers across jurisdictional boundaries. Our model pinpoints key structural parameters capturing, among other things, the strength of spillovers and of public good agglomeration economies. The key ingredient of the model is the nesting of fiscally autonomous jurisdictions in larger geographic areas within which one suspects there are spillovers, and between which one assumes zero spillovers. Using the model’s structure, one can exploit migration and house price responses to shocks in local public goods at different geographic scales— first comparing jurisdictions within nests, then comparing such nests between them— to reveal the intensity of spillovers.

We apply our framework to the French institutional setting. France boasts 35,000 fiscally

autonomous municipalities and around 2,000 municipal federations. The choice of municipal federations as nesting areas for spillovers is a natural benchmark as federations typically form geographically relevant administrative groupings of close-by municipalities. Using our rich set of databases covering the universe of French municipalities and the empirical specifications deriving from the model, we provide graphical DiD-type evidence on the impact of public good supply shocks on migration, housing consumption and housing prices. We then structurally estimate the model with GMM using moment conditions that build on these evidence. DiD evidence and GMM estimation rely on plausibly exogenous variations in government subsidies that instrument changes in the supply of local public goods. We find that public spending in a municipality accounts for 4–11% of the local public good bundle enjoyed by its residents, and that public spending in each municipality belonging to the same municipal federation accounts for an average 3.2–3.5% of this bundle.

These high estimates suggest that residents consume public goods provided by close-by municipalities almost as much as public goods available in their own municipality. The fact that French municipalities are in general very small certainly drives this result, together with the near-absence of legal restrictions — such as zoning — that would prevent cross-border public good consumption. In a simple application of our method, we simulate the impact of redefining French municipal boundaries and find strong welfare gains.

Chapter 2 was written with Benjamin Carantino and is entitled “Optimal Spatial Policies with Local Public Goods and Unobserved Location Preferences”. We further connect urban economics and public finance topics by studying the normative implications of public good agglomeration externalities and heterogeneous location preferences. Using the local public finance and population data introduced above, we present new graphical evidence suggesting substantive scale economies in the consumption of local public goods — per capita public spending decreases with population with an elasticity of 0.32 — taking due account of public assets for the first time.

We then build an economic geography model with mobile workers having heterogeneous preferences for the various locations, imperfectly elastic housing supply, endogenous public goods and tax competition between jurisdictions. In our framework, exogenous local productivity and residential amenities stand for all dimensions of physical geography. There is a central government that taxes and redistributes across places and provides a national pub-

lic good. Because the government is unable to observe location preferences, both efficient transfers that correct for public good externalities and transfers that achieve redistributive objectives are necessarily place-based and distort migration decisions. In this context, it is natural to investigate the circumstances under which Pareto-improving policies exist.

We first show that heterogeneity in location preferences increases the equity cost of migration — transfers — relatively to its efficiency benefits — agglomeration gains — so that Pareto-improving reforms only exist when preference dispersion is small enough. We then give a general sufficient statistics characterization of the transfers implementing efficient population distribution using a reduced set of variables and parameters. We derive a computable efficiency test for observed allocations and apply it to French data: ignoring location preferences leads to recommending reforms that mistakenly redistribute towards the densest and richest places. This strongly speaks in favour of better accounting for the normative implications of geographic preferences.

Finally, we carry out a revealed preference exercise by empirically investigating the structure of the social welfare weights implied by the current spatial tax and transfer system. Our results suggest that the French planner is roughly utilitarian, but that it further compensates low-density areas in a way that is orthogonal to utilitarianism. This gives support to the idea that the government designs transfers tackling horizontal inequity in access to public goods, which we argue arises from the interplay of unequal density gains from scale economies and heterogeneous location preferences.

Chapter 3 is entitled “The Deadweight Loss of Property Transaction Taxes” and provides quasi-experimental evidence on behavioral responses to property transaction taxes. I use the DV3F dataset covering all property transactions in France over the period 2010–2016 and exploit variations provided by a 2014 reform that let French counties (*départements*) increase stamp duties levied on property transactions from 5.09% to 5.81% of the purchase price.

I first document that tax changes were salient and led to significant re-timing responses which increased market activity in the average county by 76% in the week prior to the reform and depressed it by 31% in the following two months. Exploiting the phasing-in of tax increases across French counties, I implement a synthetic control method on low-level geographic treatment and control units to investigate whether increasing transaction taxes had a long-term impact on property markets. I present DiD-type graphical evidence of extensive

responses (i.e., of less transactions taking place) and estimate a net tax elasticity in the range $[0.22, 0.25]$, larger than that of any placebo reform.

I find no effect on transaction prices nor house quality, suggesting that the supply of existing homes may be very elastic, even in the short run. I illustrate an alternative mechanism in a stylized search model. A tax increase may upgrade the average match surplus to be shared between buyers and sellers, and push bargained prices up, by discouraging the search effort of buyers with lower reservation values. While increasing taxes may decrease the number of mutually beneficial transactions and lead to efficiency losses, it may improve allocative efficiency by increasing gains from trade through changes in the composition of active buyers. In this simple setup, I find that this positive composition effect likely outweighed the negative effect of lost transactions following the 2014 reform.

Chapter 1

This Town Ain't Big Enough?

Quantifying Public Good Spillovers^{*}

Despite long-standing theoretical interest, empirical attempts at investigating the appropriate level of decentralization remain scarce. This paper develops a simple and flexible framework to test for the presence of public good spillovers between fiscally autonomous jurisdictions and to investigate potential welfare gains from marginal fiscal integration. We build a quantitative spatial equilibrium model with many local jurisdictions, mobile households and endogenous local public goods causing spillovers across jurisdictional boundaries. We show how one can exploit migration and housing price responses to shocks in local public goods at different geographic scales to reveal the intensity of spillovers. Applying our framework to the particularly fragmented French institutional setting, we structurally estimate the model using a unique combination of municipal administrative panel datasets. Estimation relies on plausibly exogenous variations in government subsidies to instrument changes in the supply of public goods. We find that public goods in a municipality account for 4–11% of the local public good bundle enjoyed by its residents, and that public goods in each neighbor municipality account for an average 3.2–3.5% of this bundle. Finally, we simulate the effect of a reform increasing fiscal integration and find substantial welfare gains.

^{*}This chapter is based on a joint work with Aurélie Sotura.

1.1 Introduction

Take an economy divided into geographically distinct jurisdictions. Who should be providing local public goods? Local governments or the central government? If local jurisdictions are tasked with providing public goods, what should be their boundaries? Since the seminal works of Tiebout (1956) and Oates (1972), academics have investigated the optimal balance of power between local and more centralized forms of government. Simply put, the political economy of centralized decision-making misallocates local public services. On the other hand, decentralization may be inefficient because of spatial spillovers, i.e., the extent to which a jurisdiction's local public goods also benefit its neighbors.¹ When spillovers are strong, local jurisdictions may under-provide local public services as they do not internalize their benefits to neighbor jurisdictions. In addition, jurisdictions may actively free-ride on neighbor jurisdictions' public goods, worsening the under-provision problem.

In practice, there seems to be no consensus about the optimal size and autonomy of local jurisdictions. There remains substantial variation in institutional geography across western countries. Average town population is 4,300 in the EU compared to 16,000 in the US. Within the EU itself, there are large disparities between otherwise comparable countries. Average town density is similar in Germany (179 inhabitants per km^2), France (154 inh. per km^2) and Spain (177 inh. per km^2). However, with mean jurisdiction population of respectively 7,100 and 5,800, German and Spanish towns are much larger than the average French town only home to 1,753 inhabitants.

Depending on the magnitude of decentralization and centralization inefficiency costs, jurisdictional fragmentation may have important welfare consequences. Policy and institutional solutions to remedy spillover inefficiencies typically include Pigovian subsidies and boundary redefinition. While the former solution requires the tailoring of subsidies to spillover intensity, one simply needs knowledge of the presence of spillovers between jurisdictions belonging to a given geographic area to implement the latter. However, while a rich theoretical literature studies efficient fiscal federalism (see Oates 2005 for a review), there are few empirical attempts at structurally quantifying local public good spillovers. Indeed, there are empirical and theoretical challenges to disentangling public good spillovers from other general equilib-

¹Decentralization may be inefficient for other reasons, e.g., tax competition between local governments (see for example Brueckner 2000, Brueckner 2003 and Boadway and Tremblay 2012 for reviews). In this paper, we focus on local public good spillovers in the tradition of the early literature on decentralization. Our results are however robust to the presence of tax competition.

rium mechanisms. For example, spillovers may worsen migration externalities when public goods are subject to agglomeration or scale economies.²

Our paper revisits the local public good provision debate in a quantitative spatial equilibrium model. Our contribution is twofold. First, we develop a simple framework to test for potential welfare gains from arbitrary increases in the level of centralization of public goods provision. We ground it on a location choice model borrowed from the urban economics literature that allows us to isolate public good spillovers from other mechanisms at play in equilibrium. The key ingredient is the nesting of fiscally autonomous jurisdictions in geographic areas within which one suspects there are cross-border spillovers. By studying migration and housing price responses to local public good shocks first between jurisdictions within nests, then between such nests, one can reveal the presence of spillovers. Because these nesting areas can be made to encompass an arbitrary number of jurisdictions, our framework allows to repeatedly test for the presence of spatial spillovers until finding the jurisdictional fragmentation for which they vanish. To the best of our knowledge, this paper is the first attempt at taking a structural approach to fiscal decentralization. Second, we apply our framework to French data and provide new estimates for structural parameters that are key to the local public finance debate. In the much fragmented French context, we estimate strong spatial spillovers and public good scale economies.

Our approach has two limitations. First, assessing the inefficiency cost of centralization is beyond the scope of the present paper. These costs typically stem from the interaction between centralized political frictions that tend to create winners and losers, and heterogeneous local needs.³ As such, we abstract from the taste heterogeneity motivating the standard Tiebout literature and do not account for interactions between local and central governments. Second, structural parameters may themselves be endogenous to local public goods, e.g., to the development of regional transport.

We first develop a spatial equilibrium model that draws on the seminal framework of [Rosen \(1974\)](#) and [Roback \(1982\)](#) and allows for endogenous wages, rents and local public good provision. We let households be potentially infra-marginal in their migration choices by introducing heterogeneous preferences for jurisdictions. Agents are otherwise homogeneous in

²When public goods are not fully rival, the costs of providing residents with a given level of public good benefits increase less than one for one with population. Denser jurisdictions typically provide more public goods for less taxes, which creates an agglomeration force.

³See for example [Carbonnier et al. \(2008\)](#) for an attempt at assessing both the costs of centralization and decentralization.

skills and have identical preferences for public goods over private consumption. They vote for local taxes and a level of public goods that in turn affects equilibrium demand for jurisdictions. Most importantly, we allow for cross-border public good spillovers in a simple flexible structure that keeps the model amenable to reduced-form empirical analysis. Our model pinpoints key structural parameters related to spillovers, local public good agglomeration economies, preference for public goods, household mobility and housing supply elasticity.

Home to around 35,000 autonomous local jurisdictions accounting for 38% of EU's total, France is a natural context in which to apply our framework. We first provide difference-in-differences (DiD) evidence on the impact of public good supply shocks on migration, housing consumption and housing prices using comprehensive administrative datasets on French municipalities. We combine data covering local taxes and public spending, population, housing consumption, wages and housing prices from 2000 to 2016. Our identification strategy exploits plausibly exogenous variation in public investment subsidies coming from higher layers of government — e.g., regions or the central government — to instrument changes in local public goods. Overall, we find significant migration responses and housing price capitalization which we interpret as evidence that households are mobile, enjoy local public services and that housing supply is not inelastic in the medium run. However, we find that migration responses to public good shocks within municipal federations — which are administrative groups of close-by municipalities — are smaller than when comparing these groups between them. We also estimate significant housing price capitalization in the latter case and not in the former. In line with our theoretical framework, a candidate mechanism to explain these reduced-form results is the presence of spatial spillovers, rendering location decisions less relevant for the enjoyment of publicly provided amenities within groups of close-knit jurisdictions.

We take our theoretical framework to the data and use a Generalized Method of Moments (GMM) to estimate the model's parameters. We express moments conditions in terms of the model's structural residuals and investment subsidy shocks. We combine moment conditions related to demand for municipalities and average demand for municipal federations. Importantly, the model is identified using the same subsidy shocks that we exploit in our DiD analysis. Different variations in subsidy shocks differently affect demand for municipalities and local housing supply and demand, in particular as we allow for local housing supply elasticities to vary across municipalities.

We find significant cross-border spillovers. Our results suggest that for the average mu-

municipal federation size, public spending in a French municipality accounts for 4–11% of the bundle of local public goods enjoyed by its residents, and that public spending in its average neighbor municipality—belonging to the same municipal federation—accounts for 3.2–3.5% of this bundle. These high estimates for cross-border spillovers suggest that residents in a given municipality benefit from neighbor municipalities’ public goods almost as much as their own municipality’s. This may be because French municipalities are in general very small, and because there is almost no legal restriction such as zoning to benefiting from other municipalities’ public goods. In a simple application of our method, we simulate the impact of redefining French jurisdictions’ boundaries along pre-existing administrative lines and find strong welfare gains.

Our paper is related to the vast literature on fiscal decentralization. In his seminal paper, [Tiebout \(1956\)](#) argues that decentralized public good provision is efficient because people “vote with their feet” to choose their optimal bundle of taxes and public goods. However, [Besley \(1981\)](#) provides a formal treatment of Tiebout’s ideas and concludes that this efficiency result only holds when assuming away interesting features such as spatial spillovers and agglomeration externalities. Following [Oates \(1972\)](#), a rich theoretical literature investigates the consequences of spillovers on local public good provision and efficient federalism (e.g., [Gordon 1983](#); [Wellisch 1994](#); [Conley and Dix 1999](#); [Lockwood 2002](#); [Besley and Coate 2003](#); [Bloch and Zenginobuz 2007](#); [Cheikbossian 2008](#); [Bloch and Zenginobuz 2015](#)).

Some empirical work tests the presence of spatial spillovers. [Solé-Ollé \(2006\)](#) investigates benefits spillovers—when households enjoy public goods of neighbor jurisdictions—and congestion spillovers—when households congest such public goods—in the case of local public spending in Spain. The author finds significant evidence of both in equal magnitude. [Case et al. \(1993\)](#) offer a test when local public goods of neighbor jurisdictions are complements and conclude to the existence of spillovers.

On the methodological side, our paper relates to the canonical spatial equilibrium framework of [Rosen \(1974\)](#) and [Roback \(1982\)](#). Drawing on the seminal logit choice setup of [McFadden \(1973\)](#), this workhorse model has since been extended to account for heterogeneous mobility frictions both for households and firms (e.g., [Fajgelbaum et al. 2015](#); [Suárez Serrato and Zidar 2016](#)). We also relate to the large literature studying Tiebout type models with endogenous public good provision ([Konishi 1996](#); [Epple and Sieg 1999](#); [Brueckner 2000](#); [Bloch and Zenginobuz 2006](#)). Our work is also related to recent research in urban economics model-

ing endogenous amenities such as Ahlfeldt et al. (2015), Diamond (2016) or Fajgelbaum and Gaubert (2018). In our setup, amenities take the form of public goods and taxes that are endogenous both because of household mobility and the local voting process.

Related empirical work has investigated Tiebout type drivers of migration decisions. Early work such as Oates (1969) studies the impact of local fiscal amenities on housing price capitalization. His estimates show that property values are positively affected by public spending on schools and negatively affected by local taxes. Banzhaf and Walsh (2008) look at the impact of a particular residential amenity, air quality, on jurisdiction density using large plant openings. The authors find that location choices are environmentally motivated. Lutz (2015) estimates significant effects of lower property taxation on residential investment and housing prices, with magnitudes depending on the elasticity of housing supply.

The rest of the paper proceeds as follows. Section 1.2 provides background on French local public finance and presents some empirical regularities. In Section 1.3, we develop our spatial equilibrium model of jurisdictions with endogenous fiscal amenities. Section 1.4 describes our data. Section 1.5 presents reduced-form evidence on the impact of local public good supply shocks on different economic outcomes. In Section 1.6, we use these shocks to structurally estimate our model with GMM. Section 1.7 presents welfare implications. Section 1.8 concludes.

1.2 Institutional Background

In this section we provide background on the French local institutional context and give some historical elements on the early acknowledgement by politicians of the existence of cross-border public good spillovers and of the potential gains from coordination. We then present some stylized facts that highlight the prevalence of municipalities in France in the provision of local public goods and lay the foundation for our empirical analysis by discussing municipalities' finances.

French local institutions belong to a four-tier system. As of 2016, the territory is divided into 35,352 municipalities (*communes*), nested in 2,000 municipal federations (*intercommunalités*), henceforth MF, 100 counties (*départements*) and 13 provinces (*régions*). Following a series of decentralization laws starting in the early 1980's, France's local authorities increasingly gained autonomy regarding local public services. Each layer of local government has the autonomy to levy taxes and is in charge of providing some public goods. The 35,352 French

municipalities represent around 38% of EU's total.⁴ This large number of municipalities means that most of them are home to very few residents — 383 in the median municipality.

The French government long acknowledged that this large number of jurisdictions may be a source of inefficiency in the provision of local public goods because of unexploited economies of scale, tax competition and poor public service coordination.⁵ Central authorities hence created financial incentives to encourage municipalities to merge into larger jurisdictions (see [Leprince and Guengant 2002](#)). Local officials however, supported by their constituents, have traditionally opposed such mergers. As a result, the number of municipalities has been fairly stable over time. There were around 38,000 municipalities in the late 18th century, compared to roughly 35,000 in 2016. To bypass political obstacles to mergers, central authorities introduced the possibility for neighbor municipalities to group into municipal federations. This new tier of local government, made up of elected officials from member jurisdictions, would allow municipalities to coordinate without losing autonomy. Initially optional, being part of a municipal federation became compulsory in 2013. As a result, the share of federated municipalities jumped from 74% in 2002 to 100% in 2016, with an average 27 member municipalities per federation.

Despite the creation of this new layer, municipalities are still the largest providers of local public goods and local cooperation beyond basic services remains limited. In 2015, the local public sector — municipalities, counties and provinces — spent €229 billion, representing 9.4% of GDP. Municipalities accounted for 41% of the total, followed by counties (35%), municipal federations (14%) and provinces (11%). When looking at local infrastructure investments alone, totalling €47 billion in 2015, the relative weight of municipalities is even greater. Municipalities' investments represented 41% of the total, twice as much as counties (21%), provinces (20%) or municipal federations (18%).

Municipalities are responsible for urban planning, transport and environmental policies (19% of their budget in 2009). They are also responsible for kindergartens and primary schools (13% of the budget) which includes school creation, location decision, management and financing, i.e., everything except teachers' hiring and payroll. Finally, they are also in charge of sports, tourism and culture (17% of their budget) including the construction and maintenance of public libraries, museums, music schools, public theaters, sports and tourist

⁴See data on local administrative areas by Eurostat <https://ec.europa.eu/eurostat/fr/web/nuts/local-administrative-units>.

⁵See report by the French Senate <https://www.senat.fr/rap/r05-193/r05-1931.html>.

facilities as well as financial support to sports or cultural events and associations. Municipal federations are mostly in charge of economic development and they may cooperate on the provision of local public goods that are of the responsibility of municipalities, without any obligation.⁶

Importantly, almost all local public goods provided by municipalities are enjoyable by any household regardless of municipality of residence. This is clearly the case for urban planning, transport and environmental policies. However, it is also the case for sports or cultural facilities and events. In the case of swimming pools and public libraries for instance, there are two kinds of preferential access rules: a special fee for residents, or a special fee for residents and residents of close-by jurisdictions. While fees might differ depending on jurisdiction of residence, access can never be denied to anyone. Access to primary schools and kindergartens is subject to zoning which in principle prevents residents of a municipality to benefit from the school quality of a neighbor municipality. However, there are ways to circumvent school zoning in practice. Furthermore, small municipalities have been facing population decline for decades. In order to meet their legal obligation to provide primary and kindergarten schooling, municipalities in rural areas are allowed to form groups regarding schooling provision.⁷

To finance local public goods, municipalities have the autonomy to levy taxes.⁸ In 2015, they raised around €50 billion or 2% of GDP in direct and indirect taxes of which 33% were from the tax on resident households, 28% from the tax on property owners, 20% from the local business tax, 2% from the land tax and the remaining 17% from various small taxes (housing transaction taxes, waste management tax, etc.).⁹ As an alternative source of funding, municipalities receive €17 billion in operating subsidies to cover operating (i.e., non-investment) expenditure. These subsidies are formula-based—loosely speaking increasing with population and decreasing with mean income—and mostly coming from the central government. Finally, municipalities receive around €13 billion from other smaller sources such as various user fees. Together, these sources of funding cover 116% of municipalities' annual operating expenses (around €69 billion in 2015), the 16% surplus being invested in infrastructure.

In 2015, French municipalities invested €19 billion in local infrastructure. Funding in the

⁶They will all be in charge of water provision and waste management starting 2020.

⁷In any case, primary schools and kindergartens only account for 13% of municipalities' budget.

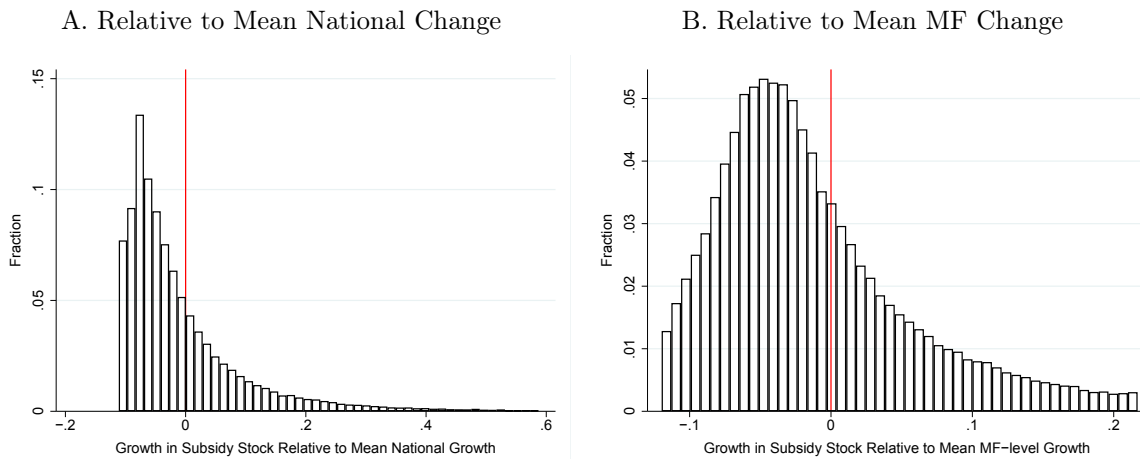
⁸This autonomy, however, is constrained by the presence of several rules limiting year-on-year variation in tax rates.

⁹Source: Ministry of Finance.

form of general endowments and investment subsidies accounted for 42% of total investment, the remaining 58% being financed by operating surpluses and additional debt. General endowments can take the form of in-kind gifts from the central government or non-targeted and automatic transfers such as VAT refunds on infrastructure expenditure. Investment-targeted subsidies however are more specific and aimed at financing well-defined investment projects. These are awarded by the boards of counties and provinces as well as by the central government to municipalities that were successful in their grant application. While we do not observe jurisdictions' applications to investment grants, we argue in Section 1.5 that these subsidies are plausibly exogenous shocks to local public good supply.

There is substantial cross-sectional variation in investment subsidies received each year. For the purpose of our paper, we define a jurisdiction's *investment subsidy stock* in year t as the sum of subsidies ever received by the jurisdiction up to t . This investment subsidy stock concept captures how much of a jurisdiction's public capital is being financed by public funds coming from a higher layer of government.¹⁰ To make yearly subsidies more comparable between municipalities, we operate some normalization. We divide the total subsidy amount

Figure 1.1: Distribution of Subsidy Stock Yearly Growth



Note: These histograms plot the distribution of the yearly growth in normalized subsidy stocks pooling years 2007, 2009 and 2010. Panel A normalizes each municipality-level observation by the national geometric mean. Panel B normalizes each municipality-level observation by the geometric mean of all municipalities belonging to the municipality's MF. In each Panel, top and bottom 1% observations are censored for exposition purposes.

¹⁰Subsidies financing durable investments are recorded as a liability stock in the municipal accounts. They are depreciated at the same speed as the investment they help financing to keep reflecting their current contribution to local assets.

received in year t by the subsidy stock in year $t - 1$. The normalized yearly subsidy can then be interpreted as the growth in the municipal subsidy stock. We then subtract to each municipal-level observation the national or own MF average. In Figure 1.1 we report the resulting distributions pooling subsidy stocks' yearly growth in 2007, 2009 and 2010.¹¹ Panel A shows the pooled distribution of this percentage change in excess of the national percentage change. It exhibits substantial variation with the 1st percentile being at -11% and the 99th percentile at $+59\%$. Panel B shows a slightly modified distribution, where each jurisdiction's subsidy percentage change is considered relatively to the mean change in the jurisdiction's MF. Again, it exhibits substantial dispersion with the 1st percentile being at -19% and the 99th percentile at $+51\%$.

1.3 Theoretical Model

This section develops our spatial equilibrium model. It extends the seminal framework of Rosen (1974) and Roback (1982) to account for heterogeneous preferences for jurisdictions in the spirit of the recent urban economics literature. Most importantly, we allow for endogenous rents, wages, local public goods and taxes. A distinguishing feature of our model is its focus on local public goods. Endogenous fiscal amenities—taxes and public spending—are central in households' location decisions and are determined through an elementary voting mechanism.

The model shares some methodological features with the urban or economic geography frameworks of Busso et al. (2013), Ahlfeldt et al. (2015), Suárez Serrato and Zidar (2016), Diamond (2016) and Fajgelbaum and Gaubert (2018). There is a finite collection of J jurisdictions indexed by j with fixed boundaries, as well as a finite collection of A mutually exclusive geographic areas indexed by a in which the J jurisdictions are nested. Because these nesting areas can be made arbitrarily large, our framework is flexible enough to accommodate many institutional settings. We note a_j the area j belongs to.

There is a continuum of imperfectly mobile households of measure 1, N_j being the share of households living in jurisdiction j . Households inelastically supply one unit of labor in their jurisdiction of residence. Jurisdiction j is characterized by a vector of endogenous observables—wage w_j , rental price r_j , aggregate public good measure \mathcal{G}_j , ad-valorem local housing tax τ_j^h and local business tax τ_j^k —as well as unobserved residential amenities. Local public

¹¹Our data are from the French municipal financial accounts (*Balance Comptable des Communes*) that we present in more details in Section 1.4.

goods are financed by local housing and business taxes as well as subsidies coming from the central government. A national proportional income tax τ^w finances government subsidies to the J jurisdictions. The sections below describe how demand for jurisdictions, housing supply, wages and local public goods are endogenously set in equilibrium.

1.3.1 Preferences

In order to easily connect theory and empirical analysis, we develop our conceptual framework in a Cobb-Douglas environment. Utility of household i living in jurisdiction j is

$$U(\mathcal{C}, \mathcal{G}, i, j) = \mathcal{C}^{1-\phi} \mathcal{G}^\phi \mathcal{E}_j^A \exp(\mu_{ij}) \quad (1.1)$$

where \mathcal{G} is the aggregate local public good and \mathcal{C} is a measure of aggregate private consumption, μ_{ij} represents the idiosyncratic taste of individual i for jurisdiction j and \mathcal{E}_j^A captures jurisdiction j 's exogenous residential amenities. Parameter $\phi \in [0, 1]$ is the taste for public goods relative to private consumption, that we assume homogeneous across households.

Consumption Agents enjoy aggregate private consumption \mathcal{C} defined by

$$\mathcal{C} = c^{1-\alpha} h^\alpha \quad (1.2)$$

where c is consumption of the nationally traded good taken as the numéraire and h is m^2 housing consumption. Parameter $\alpha \in [0, 1]$ is the housing consumption share which we assume to be constant across households. Given post-tax rental prices $r(1 + \tau^h)$ and net income $(1 - \tau^w)w$, consumption of the numéraire good is

$$c = (1 - \tau^w)w - r(1 + \tau^h)h \quad (1.3)$$

Public good The aggregate public good measure \mathcal{G}_j depends on the congested local public good of the jurisdiction households live in and on that of all other jurisdictions belonging to the same area a_j due to the presence of cross-boundary spillovers. For a jurisdiction j , we assume that congested public goods of other jurisdictions belonging to area a_j enter \mathcal{G}_j with equal spillover weights while spillovers coming from jurisdictions outside of a_j are zero. This binary structure for spillovers' spatial decay is simplistic yet allows to flexibly test for their presence in different settings. We define \mathcal{G}_j as the geometric average of j 's own congested

public good and that of other municipalities of a_j :

$$\mathcal{G}_j = \mathbb{G}_j^\delta \prod_{j' \in a_j} \mathbb{G}_{j'}^{(1-\delta)/|a_j|} \quad (1.4)$$

where \mathbb{G}_j represents the congested public good in j and $|a_j|$ is the cardinal of a_j , i.e., the number of jurisdictions in the geographic area j belongs to. Parameter $\delta \in [0, 1]$ controls the intensity of spillovers, that is, the extent to which households benefit from local public goods of adjacent jurisdictions. When $\delta = 1$ there are no spillovers and residents only enjoy the public goods provided in their jurisdiction. When $\delta = 0$ there are full spillovers within any area a so that conditional on living in a , jurisdiction of residence does not matter for the enjoyment of public goods present in jurisdictions of a . We adopt a symmetric approach for modeling congestion. We model \mathbb{G}_j , the congested public good of jurisdiction j as

$$\mathbb{G}_j = \frac{G_j}{\left(N_j^\delta \prod_{j' \in a_j} N_{j'}^{(1-\delta)/|a_j|} \right)^\kappa} \quad (1.5)$$

where G_j is the quantity of local public goods provided by jurisdiction j , which are endogenously determined along with local taxes. Parameter $\kappa \in [0, 1]$ controls the intensity of congestion. When $\kappa = 1$, public good is fully rival and public good benefits are appropriately measured by per capita spending. When $\kappa = 0$, public good is fully non-rival and public good benefits are appropriately measured by absolute public spending. As such, κ is a parameter central to agglomeration economies. The spillover parameter δ is also involved in determining the amount of congestion deteriorating the benefits from G_j . It controls how much of public service congestion is coming from adjacent jurisdictions as a direct consequence of symmetric benefit spillovers. Absent spillovers, $\delta = 1$ and public good in j is only congested by residents of j . When $\delta = 0$, public good in j is equally enjoyed and congested by all residents of a_j .

Residential amenities Jurisdiction j is further characterized by the unobserved residential amenities \mathcal{E}_j^A . They capture the mean appeal of the jurisdiction's fixed characteristics across individuals and include traditional amenities such as the weather, geographic location, etc. They also capture time-varying amenities other than those explicitly modeled. These amenities are equally valued by all residents of j .

Idiosyncratic tastes Each individual i is characterized by a vector $\{\mu_{i1}, \dots, \mu_{iJ}\}$ of idiosyncratic tastes. These μ_{ij} 's are assumed to be i.i.d. across and among individuals and distributed Extreme Value Type-I with parameters $(0, \sigma)$. They represent individual-jurisdiction specific utility premiums and notably capture heterogeneity in mobility costs and in the valuation of jurisdictions' fixed amenities.¹² Parameter σ controls the dispersion of these idiosyncratic tastes and is inversely related to household mobility. When σ is higher, density around the indifference threshold between any two jurisdictions is thinner as more households are infra-marginal. As a consequence, the migration response to a marginal change in the appeal of one jurisdiction relative to the other gets smaller.

Model parameters to be estimated so far are $\{\sigma, \phi, \kappa, \delta\}$ capturing inverse household mobility, taste for the public good, public good congestion and cross-boundary spillovers. The housing consumption share α will be calibrated from the literature.

1.3.2 Conditional Housing Demand

Conditional on living in j , agent i decides how much housing to consume while being net wage, housing price and tax taker. Given the constant share assumption, conditional individual housing demand and numéraire consumption equal

$$\begin{aligned} h_j^D &= \alpha \frac{(1 - \tau^w) w_j}{r_j (1 + \tau_j^h)} \\ c_j &= (1 - \alpha) (1 - \tau^w) w_j \end{aligned} \tag{1.6}$$

and do not depend on i . Per capita housing and numéraire consumption may hence be treated as endogenous jurisdiction amenities.

1.3.3 Demand for Jurisdictions

Agent i chooses to live in the jurisdiction that maximizes $U(\mathcal{C}, \mathcal{G}, i, j)$. We can write:

$$\ln U(\mathcal{C}, \mathcal{G}, i, j) = v_j + \mu_{ij} \tag{1.7}$$

¹²An interpretation of a higher value for μ_{ij} relative to any other $\mu_{ij'}$ is j being the jurisdiction in which i was born, educated and socialized. Another interpretation is heterogeneity in preferences for local exogenous amenities (e.g., weather, natural amenities, etc.).

where

$$v_j = (1 - \phi) \ln(\mathcal{C}_j) + \phi \ln(\mathcal{G}_j) + \ln(\mathcal{E}_j^A) \quad (1.8)$$

Households first solve for optimal housing and numéraire good consumption conditional on jurisdiction of residence according to (1.6). Based on their idiosyncratic tastes μ_{ij} , they make the extensive margin choice of where to live upon observing local aggregate private consumption $\{\mathcal{C}_j\}$, aggregate local public goods $\{\mathcal{G}_j\}$ as well as exogenous amenities $\{\mathcal{E}_j^A\}$. When comparing different jurisdictions, households are thus v_j takers. This conditional logit setup was first introduced by McFadden (1973) in a broader context of discrete choices. Demand for jurisdiction j then equals the expected set of households for which j yields the highest utility, i.e., $N_j = \mathbb{E} \left[\mathbb{1}_{\{u_{ij} > u_{iq} \forall q \neq j\}} \right]$. Because idiosyncratic tastes are distributed Extreme Value Type-I and enter utility separately from other components, population in jurisdiction j is equal to

$$N_j = \frac{\exp(v_j/\sigma)}{\sum_{j'} \exp(v_{j'}/\sigma)} \quad (1.9)$$

Loosely speaking, demand for jurisdiction j is the ratio between how attractive the jurisdiction is and the mean jurisdiction appeal in the country. Equation (1.9) notably implies that no jurisdiction is empty and that the market for jurisdictions clears:

$$\sum_j N_j = 1 \quad (1.10)$$

Total housing demand in jurisdiction j is then the result of intensive margin consumption and extensive margin in-migration:

$$H_j^D = h_j^D N_j \quad (1.11)$$

1.3.4 Housing Supply

We assume that a representative absentee landlord has the opportunity to put existing homes on the market or to develop new ones, rented at a price r per m^2 of housing. The marginal opportunity or development cost is increasing in the quantity of housing already on the market and decreasing in the jurisdiction area T . Formally, the cost of providing housing is $\left(\frac{H}{T_j}\right)^{1+\frac{1}{\eta_j}} \mathcal{E}_j^C$ where η_j is the housing supply elasticity. Further differences in local housing supply determinants are captured by the cost shifter \mathcal{E}_j^C . Since it enters the housing supply cost function in a multiplicative way, it is isomorphic to a reduction in available land. Profit

maximization yields the inverse housing supply equation

$$\ln(r_j) = \frac{1}{\eta_j} \ln\left(\frac{H_j^S}{T_j}\right) + \ln(\mathcal{E}_j^C) \quad (1.12)$$

1.3.5 Labor Demand

Local labor markets are not the focus of this paper. However, our analysis needs to account for endogenous wages as they are potentially affected by endogenous business taxation and productivity, e.g, through public investments. We make two standard assumptions. First, each jurisdiction produces a quantity Y_j of the traded good with a constant return to scale technology using local labor and capital. Second, absentee capital owners supply capital in all jurisdictions at a fixed interest rate R . We further assume that local business taxation is proportional to the outflow of local interest payments accruing to capital owners. With these assumptions, one easily shows that the inverse labor demand equation has the following form:

$$w_j = \theta_j^Y(\{G\}_{j'}) f\left((1 - \tau_j^k) R\right) \quad (1.13)$$

with w_j the local wage, τ_j^k the local business tax, f any function and $\theta_j^Y(\{G\}_{j'})$ the local labor productivity that may be endogenous to the quantity of public goods $G_{j'}$ of any jurisdiction j' in a_j . With this formulation, local wages are negatively affected by an increase in the local business tax and are positively affected by productivity shocks. Our empirical analysis would be unchanged if we allowed for tax competition and a more general production function with decreasing returns to scale (e.g., because of land use), productivity agglomeration gains or productivity spillovers as long as we econometricians observe the equilibrium wage. We would simply replace our inverse labor demand equation by a reduced-form expression such as

$$w_j = f_j\left(R, \{N, \tau^k, \theta^Y\}_{j'}\right)$$

1.3.6 Public Good Supply

As mentioned in Section 1.3.1, G_j is the quantity of public goods provided by jurisdiction j .¹³ Local taxes — ad-valorem housing tax τ_j^h and business tax τ_j^k — and subsidies F_j from higher layers of government fund the production of G_j . In our setup, F_j is exogenous. Residents of

¹³Section A.2 details how we construct G from available data. In our empirical application, G includes both the flow of public services stemming from the use of public infrastructure and the flow of local public services that are typically staff expenditure, maintenance costs, subsidies to associations, etc.

jurisdiction j choose the triplet $\{G_j, \tau_j^h, \tau_j^k\}$ that maximizes their utility subject to the municipal budget constraint. Because we assume homogeneous preferences for private consumption, the voting mechanism is akin to a maximization problem by a local social planner.

Residents' preferred policy is found by maximizing

$$(1 - \alpha) \ln \left((1 - \tau^w) w - r (1 + \tau^h) h \right) + \alpha \ln h + \frac{\phi}{1 - \phi} \delta \ln G \quad (1.14)$$

over $\{G, \tau^h, \tau^k\}$ subject to jurisdiction budget constraint

$$\Gamma_j(G) = \tau^h r H + \tau^k R K + F \quad (1.15)$$

where Γ_j maps public good quantities into the required public funds. Subsidies are financed by a national income tax τ^w that endogenously adjusts so that national budget is balanced:

$$\tau^w = \frac{\sum_j F_j}{\sum_j N_j w_j} \quad (1.16)$$

Although we do not solve for local policies in the comprehensive case, this framework is useful for our welfare application in Section 1.7.

1.3.7 Equilibrium

We close the model by imposing market clearing conditions for the traded good and housing:

$$\begin{aligned} \sum_j Y_j &= \sum_j N_j c_j \\ H_j^S &= H_j^D \quad \text{for all } j \end{aligned} \quad (1.17)$$

Definition 1. *Given the model's parameters $\{\sigma, \phi, \kappa, \delta, \alpha\}$ and $\{\eta_j, T_j\}$, local fundamentals $\{\mathcal{E}_j^A, \mathcal{E}_j^C\}$, functional forms $\{f, \Gamma_j, \theta_j^Y\}$ and national subsidies $\{F_j\}$, an equilibrium is defined by jurisdiction population $\{N_j\}$, individual consumption $\{\mathcal{C}_j, c_j, h_j\}$, housing supply $\{H_j\}$, local prices $\{w_j, r_j\}$, public goods $\{G_j, \mathbb{G}_j, \mathcal{G}_j\}$, local taxes $\{\tau_j^h, \tau_j^k\}$ and the income tax τ^w , which are determined by the following system of equations: aggregate and congested local public good definitions (1.4) and (1.5); local planner's optimization problem (1.14) and (1.15); jurisdiction utility (1.8); demand for jurisdictions (1.9); conditional housing demand (1.6); housing supply (1.12); labor demand (1.13); market clearing conditions (1.17) and national budget constraint (1.16).*

Because of agglomeration forces in the model, Definition 1 may not characterize a unique equilibrium. We argue in Section 1.6 that structural parameters can be uniquely identified nonetheless.

1.3.8 Residential Amenities

We use the equilibrium equations to express changes in unobserved residential amenities as a function of structural parameters and changes in observable endogenous variables. This step provides theoretical support to our difference-in-differences strategy in Section 1.5 and is the foundation of our GMM structural estimation in Section 1.6.

We first take the log of the demand for jurisdictions equation (1.9) and plug in jurisdiction utility equation (1.8), consumption equations (1.2) and (1.6), and public good definitions (1.4) and (1.5). We find an expression for residential amenities $\ln(\mathcal{E}_j^A)$:

$$\begin{aligned}
\ln(\mathcal{E}_j^A) &= -(1-\phi)(1-\alpha)\ln(1-\tau^w) - (1-\phi)(1-\alpha)\ln(w_j) \\
&\quad - (1-\phi)\alpha\ln(h_j) - \phi\delta\ln(G_j) - \phi(1-\delta)\frac{1}{|a_j|}\sum_{j'\in a_j}\ln(G_{j'}) \\
&\quad + (\sigma + \kappa\phi\delta^2)\ln(N_j) + \phi(1-\delta^2)\kappa\frac{1}{|a_j|}\sum_{j'\in a_j}\ln(N_{j'}) \\
&\quad + \sigma\ln\sum_{j'}\exp(v_{j'}/\sigma) + \text{constant}
\end{aligned} \tag{1.18}$$

While our model is set up in a static framework, we introduce time here as our empirical strategy in the following sections relies on time variation. In this new setup, residential amenities \mathcal{E}_{jt}^A may contain time fixed effects in addition to jurisdiction fixed effects. Using (1.18) we derive two expressions describing how residential amenities change over time. First, take a jurisdiction j nested in the larger geographic area a_j . We find an expression for how jurisdiction j 's residential amenities change over time relatively to residential amenities of all jurisdictions belonging to a_j (what we call *within-a amenity changes*). Then, consider any geographic area a . We find an expression for how the average residential amenity in a changes relatively to the average amenity in all geographic areas in the country (what we call *between-a amenity changes*).

Looking at within- a and between- a relative changes bears two advantages. First, it provides a theoretical framework for our DiD and GMM estimations that does not require to fix the utility of one jurisdiction or the denominator of (1.9).¹⁴ Second and given the assumption on the structure of spillovers, it allows us to isolate the spillover parameter δ .

¹⁴See for instance the seminal work of McFadden (1973) and more recently Diamond (2016) or Fajgelbaum et al. (2015) for examples of how this technical point is dealt with. We instead absorb the unobserved denominator of equation (1.9) by expressing residuals in relative terms.

Within- a amenity changes We subtract to equation (1.18) its within- a_j average. This differences out all terms that are common across jurisdictions of a_j , including any a_j -specific time fixed effect. We then first-difference the resulting equation which absorbs any jurisdiction fixed effect. The final expression for amenity changes is

$$\begin{aligned} \Delta \ln \overline{\mathcal{E}}_{jt}^A &= (\sigma + \kappa\phi\delta^2) \Delta \ln \overline{N}_{jt} - (1 - \alpha)(1 - \phi) \Delta \ln \overline{w}_{jt} \\ &\quad - \alpha(1 - \phi) \Delta \ln \overline{h}_{jt} - \delta\phi \Delta \ln \overline{G}_{jt} \end{aligned} \quad (1.19)$$

where

$$\overline{X}_j = \frac{X_j}{\prod_{j' \in a_j} X_{j'}^{\frac{1}{|a_j|}}}$$

for any variable X and Δ is the first-difference operator between any two arbitrary periods. Note that with full spillovers ($\delta = 0$), within- a relative public good changes have no effect on within- a relative migration responses as location does not matter for the enjoyment of public goods conditional on living in a .¹⁵

Between- a amenity changes Going back to equation (1.18), we take its average across all jurisdictions of any geographic area a . Treating a as the new level of observation, we then subtract the average across all a 's in the country. This step absorbs all fixed effects common across a 's, including common time fixed effect. Alternatively, one may subtract the average across all a 's within smaller geographic units in which the a 's are nested to better account for spatial variation in the time fixed effects. We then first-difference the resulting equation which absorbs any a fixed effect. The final expression is

$$\begin{aligned} \Delta \ln \widehat{\mathcal{E}}_{at}^A &= (\sigma + \kappa\phi) \Delta \ln \widehat{N}_{at} - (1 - \alpha)(1 - \phi) \Delta \ln \widehat{w}_{at} \\ &\quad - \alpha(1 - \phi) \Delta \ln \widehat{h}_{at} - \phi \Delta \ln \widehat{G}_{at} \end{aligned} \quad (1.20)$$

¹⁵Alternatively, one can plug in the housing consumption expression from equation (1.6):

$$\begin{aligned} \Delta \ln \overline{\mathcal{E}}_{jt}^A &= (\sigma + \kappa\phi\delta^2) \Delta \ln \overline{N}_{jt} - (1 - \phi) \Delta \ln \overline{w}_{jt} \\ &\quad + \alpha(1 - \phi) \Delta \ln \overline{r}_{jt} + \alpha(1 - \phi) \Delta \ln \overline{\mathcal{T}}_{jt} - \delta\phi \Delta \ln \overline{G}_{jt} \end{aligned}$$

where $\mathcal{T} = 1 + \tau^h$.

where

$$\widehat{X}_a = \frac{\prod_{j \in a} X_j^{\frac{1}{|a|}}}{\prod_{a'} \left(\prod_{j \in a'} X_j^{\frac{1}{|a'|}} \right)^{\frac{1}{A}}}$$

for any variable X and A is the number of areas a . Importantly, these between- a amenity changes do not feature parameter δ . Indeed, public good spillovers are contained within each a while equation (1.20) is at the a level.¹⁶

1.4 Data

We take advantage of a unique combination of panel administrative datasets that we obtained from the French Ministry of Finance and Ministry for the Environment as well as publicly available datasets at the jurisdiction level, overall covering the period 2000–2016. In this section we introduce the data and describe how we construct the variables present in the model.

Sample Our sample is the universe of mainland French municipalities that experience no boundary changes between 1999 and 2016. Not dropping municipalities experiencing boundary changes — such as municipalities merging or acquiring land from others — may lead to artificial variation in their supply of local public goods, population and economic outcomes. Fortunately, very few jurisdictions experience such changes and this manipulation leaves us with 34,835 jurisdictions, i.e., 96% of them, and an almost-complete partition of the French mainland territory.

Municipal financial accounts We use detailed municipal financial accounts (*Balance Comptable des Communes*) obtained from the French Ministry of Finance for every year between 2002 and 2016. Financial accounts contain information on municipalities' detailed asset and liability position as well as the composition of their yearly budget. Although it would be of great interest to disentangle the various components of local public goods, we are unable to

¹⁶Alternatively, one can plug in the housing consumption expression from equation (1.6):

$$\begin{aligned} \Delta \ln \widehat{\mathcal{E}}_{at}^A &= (\sigma + \kappa\phi) \Delta \ln \widehat{N}_{at} - (1 - \phi) \Delta \ln \widehat{w}_{at} \\ &\quad + \alpha(1 - \phi) \Delta \ln \widehat{r}_{at} + \alpha(1 - \phi) \Delta \ln \widehat{\mathcal{T}}_{at} - \phi \Delta \ln \widehat{G}_{at} \end{aligned}$$

classify the various items by their nature as the existing categories follow accounting definitions. This is of little consequence, however, as this paper is more general and seeks to address the broad inefficiencies in the provision of local public goods and not the effect of a particular type of public good.

We construct the public good quantity G from municipal accounts by assuming that G is a Cobb-Douglas public good index capturing both durable facilities and operating expenditure that benefit residents (see Appendix A.2 for detailed methodology). We construct durable facilities as the sum of all public assets minus the raw value of the land and financial assets such as cash.¹⁷ As mentioned in Section 1.2, investments notably include schools, transportation infrastructure, parks improvements, sports facilities, museums, art collections, etc. They also include investment subsidies to other parties such as local clubs and associations. Importantly, they do not contain social housing units.¹⁸ We construct operating expenditure as the sum of staff expenditure, maintenance spending, payments for external services and operating subsidies to third parties.¹⁹

In our model, local public budgets are partly financed by public subsidies F_j coming from the central government and higher layers of local government. In our empirical application, we instrument local public goods using a subset of these subsidies, namely investment-targeted subsidies, that we call S_j . We define S_j as the sum of all investment subsidies ever received by municipality j from counties, provinces and the central government minus all associated depreciation. In 2016, investment subsidies represent 23% of the financing of public capital. Endowments account for 28% while the remaining 41% are from local contributions—past (reserves) or future (debt). As mentioned in Section 1.2, S_j depreciates together with the public capital it helps financing. We will see that S_j proves a good basis for instrumenting public goods G_j .

Population and housing We use FILOCOM (*Fichier des Logements à la Commune*) which is an exhaustive database on household housing stock. It provides information on each non-commercial dwelling every two years between 1994 and 2014 such as its location, surface,

¹⁷Taking out the raw value of the land seems natural as residents are unlikely to value it. Given that land is mostly a gift from the central government, erasing this asset as well as the corresponding liability from the balance sheet is neutral in our analysis. Cash and other liquid assets can be considered negative debt and are accounted for in our theoretical framework in the form of future taxes. They are recorded at book value and account for investment depreciation.

¹⁸Social housing units, when publicly owned, are held by ad hoc entities and not by jurisdictions.

¹⁹We exclude interests payments as they do not correspond to consumable services.

vacancy status, whether it is a main or a secondary home, rented or owner-occupied, and its number of residents. We use FILOCOM to construct our municipality-level population and housing database which contains the stock of rented or owner-occupied main homes per jurisdiction and the total and per capita housing m^2 surface. Population N_j from the model is all inhabitants of housing units for which the head of household is aged 20–65 in jurisdiction j . Total housing consumption H_j is how much m^2 are consumed by the N_j residents, and per capita housing consumption is $h_j = H_j/N_j$.

Housing prices We use data on housing prices instead of rents as there is no exhaustive database on rents at the jurisdiction level prior to 2015. We combine housing transactions database from the notary offices (named BIEN for the Parisian region and PERVAL for the rest of France). We construct a database on housing prices per m^2 for every two years between 2000 and 2014 at the municipality level. We assume that housing transaction prices from the notary data are the net present value of unobserved rents r . Proportional changes in r thus equal proportional changes in housing prices. This method bears two caveats, one methodological and the other regarding the quality of the price data. First, although tenants pay rent every year—explicitly or implicitly—not all jurisdictions record housing transactions every year. As such, we can infer the rent growth of a jurisdiction only when it experiences a transaction in two consecutive periods. Second, the coverage of housing transactions in the notary data is not exhaustive.

Local tax data We use detailed local tax data (*Recensement des Éléments d’Imposition à la Fiscalité Directe Locale*—REI) for every year from 2002 to 2016. REI features all tax bases and rates at the jurisdiction level. We compute the ad-valorem local residence tax of the model from observed residence tax revenues Rev_h , total housing stock H and housing prices p at the jurisdiction level: $\tau^h = \frac{\text{Rev}_h}{Hr}$ where housing prices are the net present value of rents $r = \frac{R}{1+R} \times p$.

Wages We use labor income data from the IRCOM (*Impôt sur le Revenu par Commune*) dataset. IRCOM summarizes labor and social security total income at the municipal level from 2002 to 2016. We construct our jurisdiction wage measure w as total labor income divided by the number of tax units reporting positive labor income.

Socio-demographics and geographic data We use census data to get jurisdiction-level information on total population and socio-demographic characteristics for years 1990, 1999, 2008 and 2013. We also use publicly provided data from the National Statistical Institute (INSEE) on jurisdictions' geography (municipal federation they belong to, distance to center of urban area center, area, etc.). Although our analysis does not study the public goods provided by municipal federations, we use these geographic groupings for the definition of the a 's. From 2000 to 2016, municipalities gradually joined MFs. In 2016, all municipalities belonged to municipal federations. We assign each municipality to its 2016 MF for the whole duration of our panel so that these geographic groupings are constant over time.

Table 1.1 gives elementary descriptive statistics on jurisdictions summarizing some of the above constructed variables.

Table 1.1: Descriptive Statistics on Jurisdictions

Variable	Mean	St. Dev.
Current spending ^a	1,517,088	23,022,598
Current spending per adult ^b	1,111	1,786
Investment stock ^a	9,383,611	98,714,400
Investment stock per adult ^b	13,481	24,130
Population ^c	1,746	14,614
Population 20–65 ^d	994	4,573
Housing price per square meter ^e	1,570	978
Net income per adult ^d	10,848	3,614
Housing service per jurisdiction ^d	16,247	83,086
Housing surface per adult ^d	39	6

Note: This table gives the averages and standard deviations across jurisdictions for some of Section 1.5 constructed variables in 2009 or the closest available year.

^a2009 municipal accounts, author's calculations

^b2009 municipal accounts, 2008 Census data, author's calculations

^b2009 municipal accounts, 2009 FILOCOM, author's calculations

^d2009 FILOCOM, author's calculations

^c2008 Census data, author's calculations

^e2008 Notaries databases, author's calculations

1.5 Reduced-Form Evidence

This section builds on the theoretical framework set out in Section 1.3 and presents the results of DiD models looking at the impact of changes in local public goods on changes in a range

of economic outcomes — population, housing consumption, housing prices and wages.

In our empirical application, the a 's are the municipal federations in their 2016 form introduced in Section 1.2. Investigating whether these geographic areas are relevant for further fiscal integration makes economic and historical sense. Indeed, as argued in Section 1.2, jurisdictions historically grouped into MFs to partly internalize public good externalities and rationalize costs. We expect a relatively higher public good supply shock to increase relative migration towards targeted jurisdictions, increase total housing consumption, bid up rents in the housing market and consequently lower per capita housing m^2 consumption. Because of public good spillovers within MFs, we expect within-MF changes in public good supply to have a smaller impact than between-MF changes. Although labor markets are not the focus of this paper, we can conjecture that potential productivity increases would increase wages.²⁰

1.5.1 DiD Framework

Changes in public goods are likely endogenous to changes in the other variables of our model. We instrument changes in local public goods using the large variation in investment-targeted subsidies for which we give descriptive evidence in Sections 1.2 and 1.4. This identification strategy would be undermined if investment subsidies were correlated with changes in jurisdictions' characteristics. Such correlation would arise, for instance, if investment subsidy were systematically granted based on jurisdictions' population, housing or labor market outcomes. We follow [Suárez Serrato and Zidar \(2016\)](#) and [Fuest et al. \(2018\)](#) and use DiD models to look at cumulative changes in a range of economic outcomes around investment subsidy shocks while relying on the absence of pre-trends to make the case for the validity of our instrument.

In line with our theoretical framework, we look at both within-MF and between-MF variations in public goods and other endogenous variables. We instrument within-MF changes in public goods with within-MF yearly changes in jurisdictions' subsidy stock $\Delta \ln \bar{S}_j$. Similarly, we instrument between-MF changes in public goods with between-MF yearly changes in jurisdictions' subsidy stock $\Delta \ln \widehat{S}_a$.

We pool shocks happening in 2007, 2009 and 2010 and keep a $[-6, 4]$ distance windows around shock years. We limit ourselves to these shocks for two reasons. First, it allows us to have a balanced panel in terms of distance to the shocks and sufficient temporal depth — i.e., 6 years pre-shock — to inspect the pre-trends. This mechanically reduces the scope to shocks

²⁰Note that although it is not modeled here, a public good supply shock may come with a public labor demand shock that would increase wages.

happening around the middle of our sample. Second, we do not include the 2008 subsidies since 2008 is a municipal election year and we suspect (and empirically confirm) that subsidies that year were much more endogenous to jurisdiction conditions.

For both within-MF and between-MF regressions, we investigate responses from population, housing per capita, total housing, wage, rents and the public good index.

Within-MF regressions We first run the following regressions:

$$\Delta \ln \bar{X}_{j,d,-6} = \beta_d^W \Delta \ln \bar{S}_{j,0,-1} + \sum_{k=-5}^{-1} \beta_k^W \Delta \ln \bar{S}_{j,k,k-1} + \mathcal{I}_j \cdot \gamma_d^W + u_{j,d}^W \quad (1.21)$$

where X is any of the model's endogenous variables. We run a separate regression for each $d \in \{-6, -4, -2, 0, 2, 4\}$ where $d = 0$ is any shock year, i.e., 2007, 2009 or 2010. The main explanatory variable is the one-year relative subsidy change $\Delta \ln \bar{S}_{j,0,-1}$. \mathcal{I}_j is a vector of flexible dummies for 1999 baseline characteristics included to control for those changes in unobserved amenities, housing supply or productivity determinants that are correlated with these observed characteristics.²¹ To deal with potential serial correlation of the instrument we also control for $\Delta \ln \bar{S}_{j,k,k-1}$ with $k \leq -1$, that is, all subsidy shocks happening before the shock year. We cluster standard errors at the MF level to account for spatial correlation of the error term as well as auto-correlation since our regressions pool shocks happening in different years.

While the instrument is the one-year log difference in our subsidy stock measure, we look at long differences in post-shock outcomes to account for construction delays, frictions in year-on-year responses by individuals and more generally for dynamic adjustments to the shock. Regression coefficients we obtain are interpreted as cumulative elasticities.

Between-MF regressions Similarly, we run the following regressions:

$$\Delta \ln \widehat{X}_{a,d,-6} = \beta_d^B \Delta \ln \widehat{S}_{a,0,-1} + \sum_{k=-5}^{-1} \beta_k^B \Delta \ln \widehat{S}_{a,k,k-1} + \mathcal{I}_a \cdot \gamma_d^B + u_{a,d}^B \quad (1.22)$$

The reduced-form elasticities β_d 's represent how the subsidy shock in period 0 affects cumulative outcome growth between -6 and $d \in [-4, 4]$. The absence of pre-trends, i.e.,

²¹Controls include normalized baseline 1999 jurisdiction population and density.

$\beta_d = 0$ for $d < 0$ for all outcomes of our within-MF and between-MF regressions would strongly support the validity of our instrument.

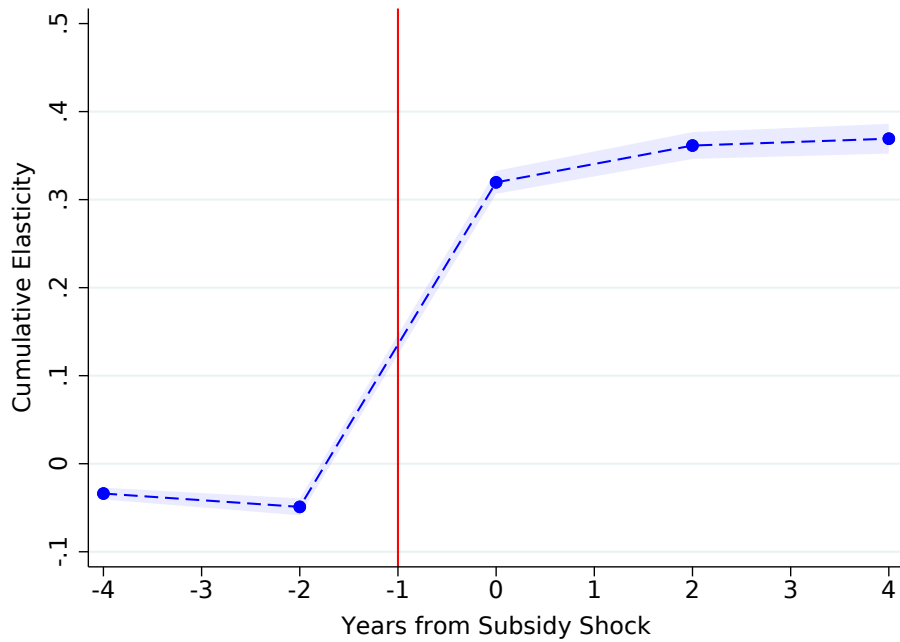
1.5.2 Results

We present graphical evidence on the evolution of our different outcomes around subsidy shocks based on our within-MF and between-MF DiD models. These graphs have no immediate “treatment effect” interpretation but show how the gradient of different outcomes is affected by presumably exogenous changes in subsidy stocks. Their contribution, however, is twofold. They provide convincing evidence of absence of selection into treatment, as well as of significant behavioral responses to changes in local public goods. They also give support to our theoretical framework as we find stronger behavioral responses — in terms of migration, housing consumption and housing prices — in the between-MF setup than in the within-MF setup, which we interpret as evidence of stronger public good spillovers within municipal federations than between.

Within-MF behavioural responses Figure 1.2 shows the first stage result, i.e., how the cumulative growth of our public good index evolves around a sudden investment subsidy shock. For each municipality, changes are relative to the mean change in the MF it belongs to. The graph reports our DiD coefficients β_d^W , that is, the effect of the shock in $d = 0$ on public good’s cumulative growth starting in $d = -6$. Visual inspection of the pre-trends confirms the exclusion restriction. Before $d = 0$, cumulative growth is flat and while it is significantly different from zero, it is not economically significant compared to the subsequent hike. After $d = 0$, growth in G becomes strongly positively correlated with the shock. A 1% change in the stock of investment-targeted subsidies leads to a 0.37% change in the public good index after five years. This effect is significant at the 1% level.

We now turn to the reduced-form effect of subsidy shocks on population. Figure 1.3 shows that treatment intensity is not correlated with municipality migration dynamics prior to $d = 0$. However, municipalities which received a relatively higher subsidy shock subsequently experienced relatively higher in-migration. A 1% relative increase in investment-targeted subsidies leads to a 0.03% relative increase in population after five years. This effect is significant at the 1% level. Put together with the first stage result, the reduced-form elasticity of population to public spending within a municipal federation is approximately 0.07. In other words, when municipality public spending grows 10 percentage points more than (geometric) mean

Figure 1.2: Within-MF Public Good Changes

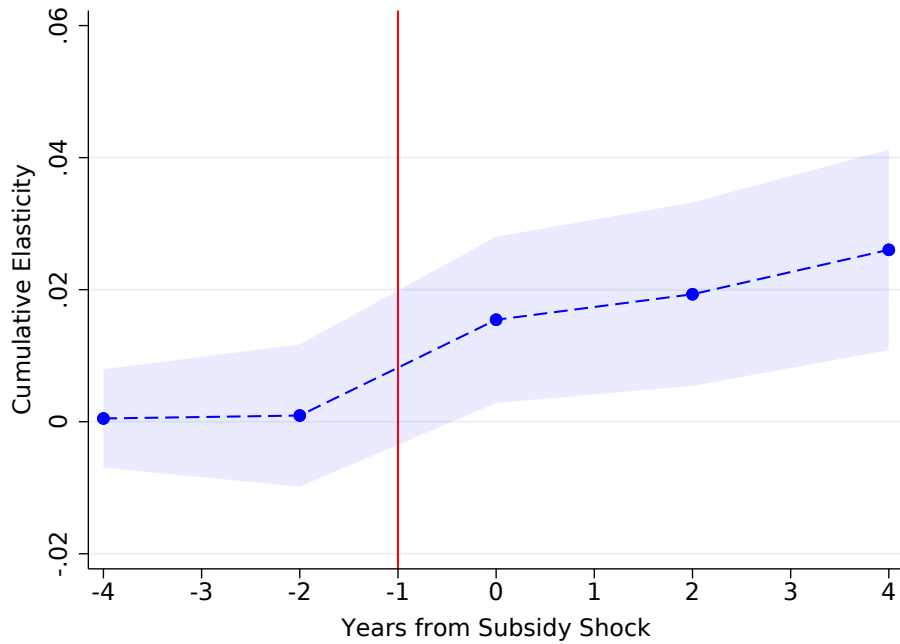


Note: This graph plots the coefficients β_d^W of Section 1.5 regressions. It shows the effect of changes in a municipality's subsidy stock in year 0 on cumulative changes in local public goods starting 6 years prior to the shock year. All changes are relative to mean changes in the MF. Standard errors are clustered at the MF level. We report the 5% confidence bands.

public spending in the MF, municipality population grows 0.7 percentage points more than (geometric) mean population in the MF.

It is useful to give a concrete example illustrating the intensity of migration responses within municipal federations. Consider a municipal federation a in its final 2016 form with mean number of member municipalities (27 municipality members) all having 2009 mean population (1,000 residents), mean investment level (€9,400,000) and mean operating expenditure (€1,540,000). Let us abstract from public good depreciation for simplicity. Assume that municipality j of a raises its investment level and yearly expenditure by 10%, that is, invests €940,000 in durable infrastructure and commits to increase yearly operating expenditure by €154,000. All other member municipalities of a decrease them by $10\%/26 = 0.38\% \approx 0\%$. Other municipalities in other MFs do not change their policies. Mean public good growth in a is zero so that a will experience no in-migration according to the model's equations. Municipality j experiences a 10 percentage points growth in public services in excess of the zero mean

Figure 1.3: Within-MF Migration Response

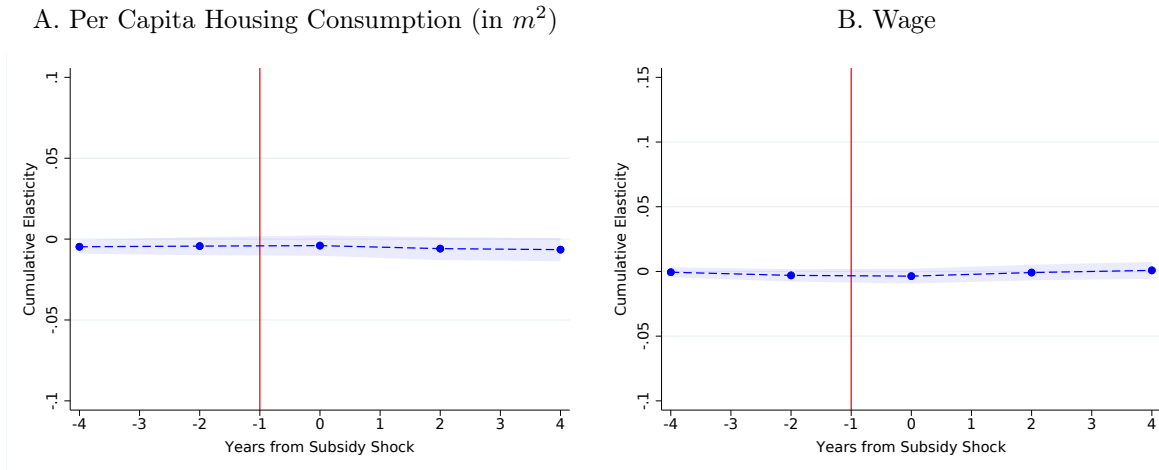


Note: This graph plots the coefficients β_d^W of Section 1.5 regressions. It shows the effect of changes in a municipality's subsidy stock in year 0 on cumulative changes in the number of residents aged 20–65 starting 6 years prior to the shock year. All changes are relative to mean changes in the MF. Standard errors are clustered at the MF level. We report the 5% confidence bands.

MF growth, and will experience a 0.7 percentage points population growth. This represents in-migration of 7 additional residents coming from out-migration from other municipalities of a for an initial investment of €940,000 and an increase of €154,000 in yearly expenditure.

Figure 1.4 shows how housing consumption per capita and wages evolve around the subsidy shock. Pre-trends are flat and not significantly different from zero in all Panels. They make an even stronger case for investment subsidies as exogenous shocks to local public good supply. None of the endogenous outcomes—public goods, population, housing consumption and wages—exhibit pre-shock dynamics correlated with shock intensity. Panel A offers evidence that subsidy shocks are not correlated with increasing or decreasing pressure in local housing markets. Furthermore, we find that per capita housing consumption, which subsumes housing prices and residence tax responses, is not affected by the subsidy shocks. Panel B gives comfort that public investment shocks are not driven by favourable or unfavourable trends in local productivity. Wages seem to be unaffected by subsidy shocks. The distance window we consider

Figure 1.4: Within-MF Per Capita Housing and Wage Responses



Note: These graphs plot the coefficients β_d^W of Section 1.5 regressions. They show the effect of changes in a municipality's subsidy stock in year 0 on cumulative changes in per capita housing consumption and wages starting 6 years prior to the shock year. All changes are relative to mean changes in the MF. Standard errors are clustered at the MF level. We report the 5% confidence bands.

might be too small to dissipate frictions in wage adjustments, however.²² Within-MF subsidy shocks hence seem to induce migration responses through changes in local public goods that do not capture significant changes in other endogenous municipality characteristics.

Table 1.2 reports alternative estimates without the different control variables. Our preferred estimates in column (3) correspond to the endpoints of Figures 1.2, 1.3 and 1.4. They are largely unaffected when we experiment with different specifications.

Between-MF behavioural responses We now investigate how MF-level changes in investment subsidies affect changes in MF-level economic outcomes. Figure 1.5 is the symmetric of Figure 1.2 and shows the first stage of our between-MF regressions. All changes are relative to mean regional changes, taking the geometric average across each municipal federation as the new level of observation. Public good cumulative growth is not significantly different from zero before the shock but it adjusts sharply after the shock. A 1% increase in the subsidy stock leads to a 0.42% increase in the public good index after five years. This effect is significant at the 1% level.

Figure 1.6 shows that MFs which received higher subsidy shocks also experienced relatively

²²We also abstract from the possibility that agents commute to neighbor municipalities to work, which would dilute effects on wages even further.

Table 1.2: Within-MF Sensitivity Analysis

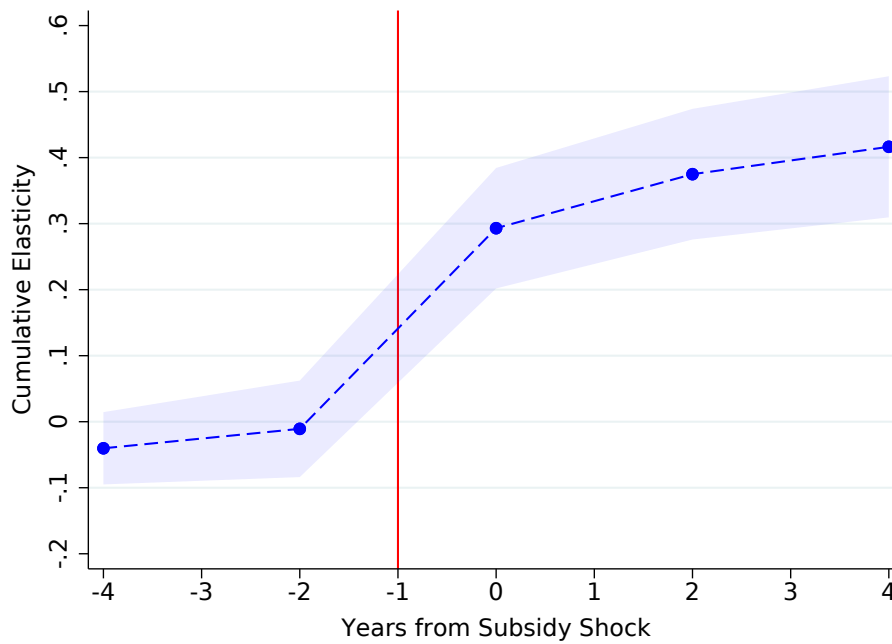
	(1)	(2)	(3)
	$\Delta \ln \bar{G}_j$	$\Delta \ln \bar{G}_j$	$\Delta \ln \bar{G}_j$
$\Delta \ln \bar{S}_j$	0.363*** (0.009)	0.372*** (0.009)	0.369*** (0.009)
	$\Delta \ln \bar{N}_j$	$\Delta \ln \bar{N}_j$	$\Delta \ln \bar{N}_j$
$\Delta \ln \bar{S}_j$	0.033*** (0.007)	0.032*** (0.008)	0.026*** (0.008)
	$\Delta \ln \bar{w}_j$	$\Delta \ln \bar{w}_j$	$\Delta \ln \bar{w}_j$
$\Delta \ln \bar{S}_j$	0.002 (0.003)	0.003 (0.003)	0.001 (0.003)
	$\Delta \ln \bar{h}_j$	$\Delta \ln \bar{h}_j$	$\Delta \ln \bar{h}_j$
$\Delta \ln \bar{S}_j$	-0.006 (0.003)	-0.006 (0.004)	-0.007 (0.004)
	$\Delta \ln \bar{T}_j$	$\Delta \ln \bar{T}_j$	$\Delta \ln \bar{T}_j$
$\Delta \ln \bar{S}_j$	0.002 (0.003)	0.002 (0.003)	0.001 (0.003)
Lag $\Delta \ln \bar{S}_j$		Yes	Yes
Baseline Controls			Yes
Observations	99,593	99,593	99,593

Note: This Table reports estimates for $\beta_{d=4}^W$, i.e., the effect of the shock happening in $d = 0$ on cumulative outcome growth between $d = -6$ and $d = 4$. Lag shocks include all past shocks up to $d = -1$. We flexibly control for baseline 1999 jurisdiction population and density. Standard errors are clustered at the MF level. * significant at 10%; ** significant at 5%; *** significant at 1%.

higher in-migration. A 1% relative shock leads to a 0.15% relative population increase after five years. This effect is significant at the 1% level. It represents a population elasticity with respect to public goods of approximately 0.37, that is, a 10 percentage points increase in public spending in excess of regional public spending growth is met with a subsequent 3.7 percentage points excess population growth. This estimate is roughly five times higher than the point estimate of within-MF regressions. As outlined in our theoretical framework, a candidate mechanism to explain this discrepancy is the presence of cross-boundary spillovers. Indeed, in the polar case of full spillovers, changes in municipality local public goods relative to mean changes in the MF should not affect within-MF migration. However, changes between MFs would still affect migration decisions as we assume spillovers abruptly die out at the MF frontier.

Again it is useful to illustrate the intensity of migration responses between municipal federations. Consider the same setting as the above example, but this time assume that *all*

Figure 1.5: Between-MF Public Good Changes

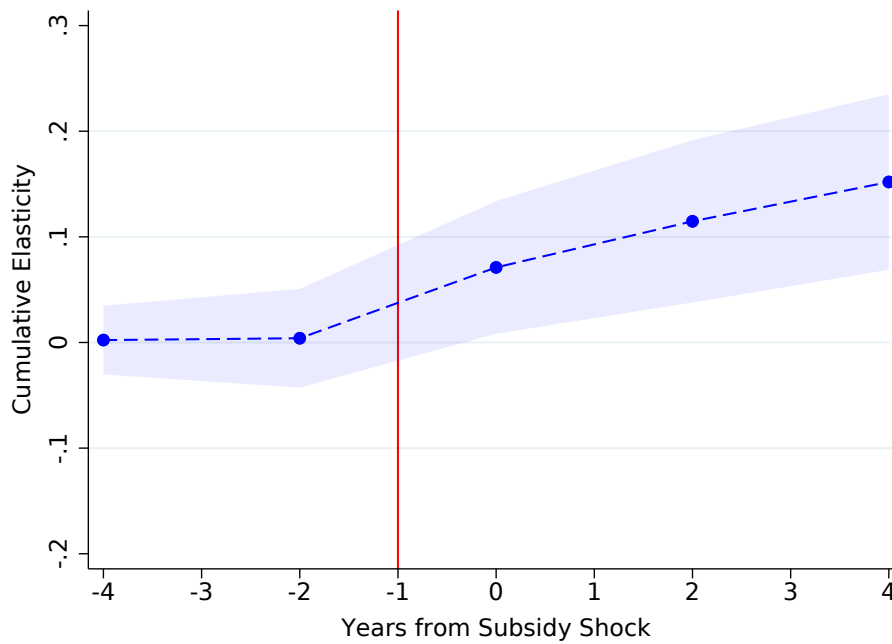


Note: This graph plots the coefficients β_a^B of Section 1.5 regressions. It shows the effect of changes in a MF subsidy stock in year 0 on cumulative changes in local public goods starting 6 years prior to the shock year. All changes are relative to regional mean changes. Standard errors are clustered at the MF level. We report the 5% confidence bands.

municipalities of a increase their investment and yearly expenditure by 10%, that is, invest €940,000 in durable infrastructure and commit to increase yearly expenditure by €154,000. All other municipalities in all other municipal federations decrease them by $10\%/1,266 \approx 0\%$. Municipalities of a experience a mean 10 percentage points growth in public services in excess of the zero mean growth, and will experience a mean 3.7 percentage points population growth. This represents in-migration of 37 additional residents coming to each municipality of a from out-migration from other federations for an initial investment of €940,000 and an increase of €154,000 in yearly services.

Figure 1.7 suggests that this reduced-form elasticity differs from the micro (i.e., absent general equilibrium adjustments) migration responses to public good changes since it does not hold constant adjustments in other local amenities. Panel A shows again that subsidies were not awarded according to local trends in housing supply determinants. However, consumable housing per capita decreases in equilibrium in response to migration pressure and increasing

Figure 1.6: Between-MF Migration Response

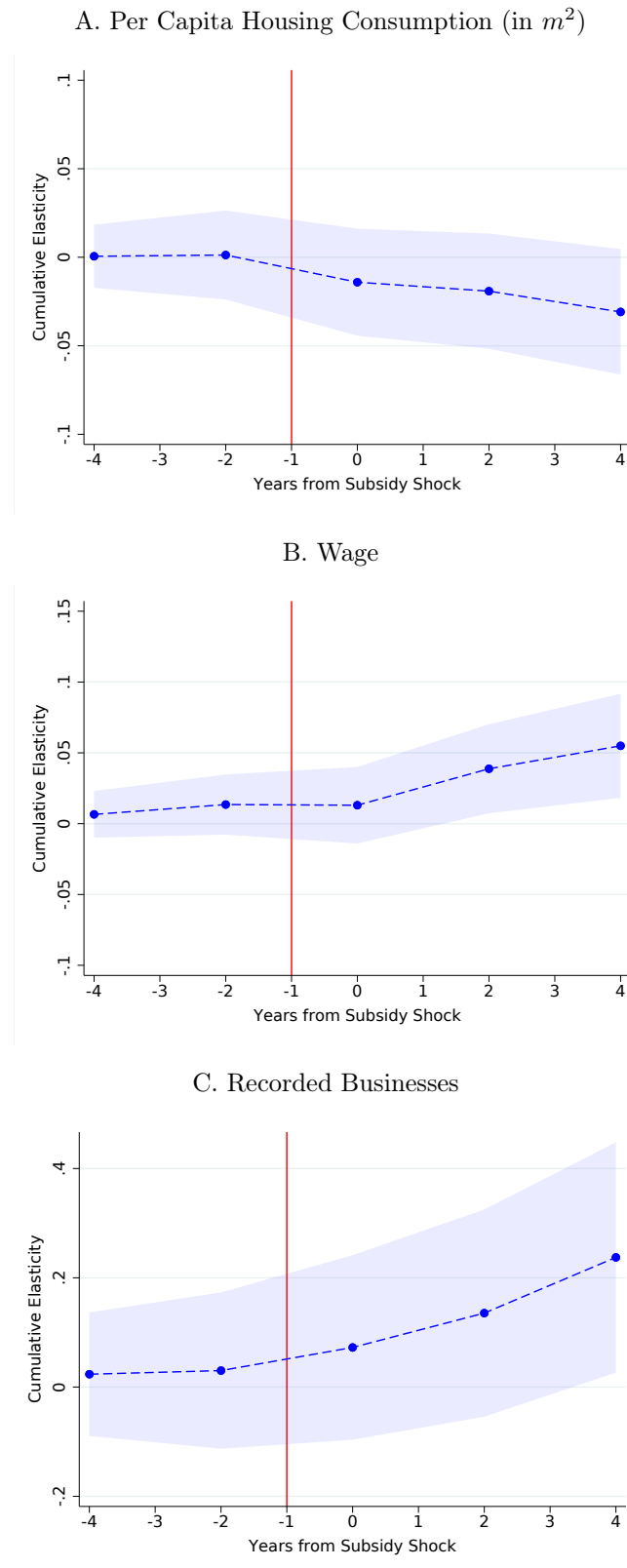


Note: This graph plots the coefficients β_d^B of Section 1.5 regressions. It shows the effect of changes in a MF subsidy stock in year 0 on cumulative changes in the number of residents aged 20–65 starting 6 years prior to the shock year. All changes are relative to regional mean changes. Standard errors are clustered at the MF level. We report the 5% confidence bands.

marginal housing supply costs with a reduced-form elasticity of housing per capita with respect to public goods of -0.07 . This also contrasts with our within-MF analysis. Panel B shows again that local productivity dynamics are not correlated with subsidy shocks. However, we find this time that wages are affected by local public goods. This result suggests that local public investment is also shifting the labor demand curve upwards by boosting local productivity. This is further evidenced by Panel C where we report the cumulative elasticities of the (geometric) average number of businesses with respect to the subsidy stock. Contrasting this result with the absence of within-MF wage variation hints at the presence of production spillovers or commuting within municipal federations. In any case, these effects are fully captured by the residential wage variation and do not bias our spillover estimates as long as we separately instrument wage changes in the GMM estimation.

Overall these results are consistent with weaker public good spillovers between municipal federations than within. Table 1.3 reports alternative estimates where we let the fixed effect

Figure 1.7: Between-MF Per Capita Housing and Wage Responses



Note: These graphs plot the coefficients β_a^B of Section 1.5 regressions. They show the effect of changes in a MF subsidy stock in year 0 on cumulative changes in per capita housing consumption, wages and number of businesses starting 6 years prior to the shock year. All changes are relative to regional mean changes. Standard errors are clustered at the MF level. We report the 5% confidence bands.

Table 1.3: Between-MF Sensitivity Analysis

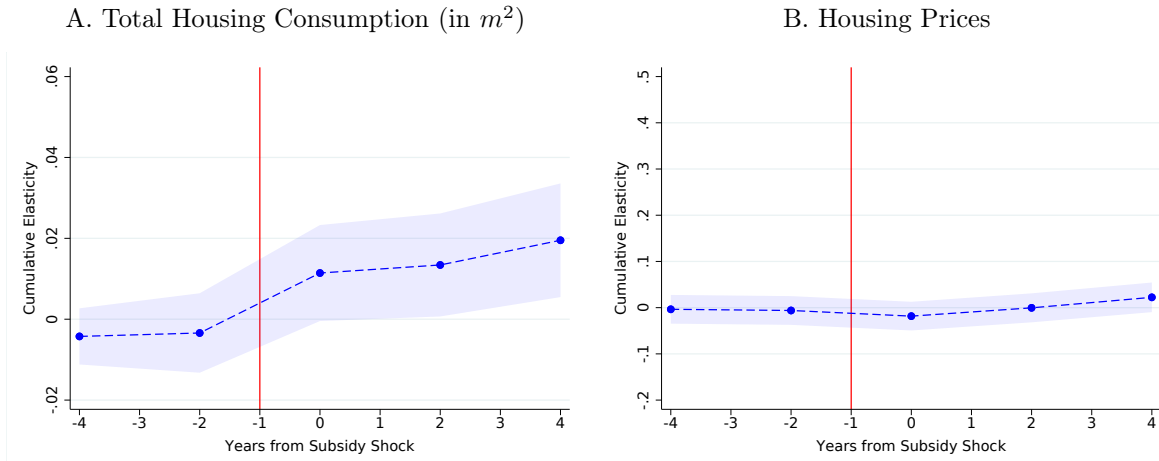
	(1)	(2)	(3)
	$\Delta \ln \widehat{G}_a$	$\Delta \ln \widehat{G}_a$	$\Delta \ln \widehat{G}_a$
$\Delta \ln \widehat{S}_a$	0.258*** (0.048)	0.416*** (0.055)	0.466*** (0.055)
	$\Delta \ln \widehat{N}_a$	$\Delta \ln \widehat{N}_a$	$\Delta \ln \widehat{N}_a$
$\Delta \ln \widehat{S}_a$	0.094** (0.038)	0.152*** (0.042)	0.148*** (0.045)
	$\Delta \ln \widehat{w}_a$	$\Delta \ln \widehat{w}_a$	$\Delta \ln \widehat{w}_a$
$\Delta \ln \widehat{S}_a$	0.008 (0.017)	0.055*** (0.019)	0.050** (0.020)
	$\Delta \ln \widehat{h}_a$	$\Delta \ln \widehat{h}_a$	$\Delta \ln \widehat{h}_a$
$\Delta \ln \widehat{S}_a$	-0.036** (0.016)	-0.031 (0.018)	-0.030 (0.019)
	$\Delta \ln \widehat{\mathcal{T}}_a$	$\Delta \ln \widehat{\mathcal{T}}_a$	$\Delta \ln \widehat{\mathcal{T}}_a$
$\Delta \ln \widehat{S}_a$	-0.017 (0.015)	-0.019 (0.016)	-0.023 (0.016)
Lag $\Delta \ln \widehat{S}_a$	Yes	Yes	Yes
Baseline Controls	Yes	Yes	Yes
Time FE X	County	Region	Nation
Observations	5,428	5,428	5,428

Note: This Table reports estimates for $\beta_{d=4}^B$, i.e., the effect of the shock happening in $d = 0$ on cumulative outcome growth between $d = -6$ and $d = 4$. Lag shocks include all past shocks up to $d = -1$. We flexibly control for baseline 1999 jurisdiction population and density. In the model, time fixed effects in the residuals are assumed to be uniform at the national level and we divide endogenous variables by their national geometric mean. Here, we allow for division by either national, regional or county geometric mean to account for year, year X region or year X county fixed effects in the structural residuals. Standard errors are clustered at the MF level. * significant at 10%; ** significant at 5%; *** significant at 1%.

in the residential amenities be at the national or county level, i.e., where we divide each MF-level observation by the national or county geometric mean instead of the regional mean. They are largely unaffected when we experiment with different specifications. While the migration response seems to be smaller when absorbing county fixed effects instead of national or regional ones, it mirrors a weaker first stage intensity so that the reduced-form elasticities are comparable. Our preferred estimates in column (2) correspond to the endpoints of Figures 1.5, 1.6 and 1.7.

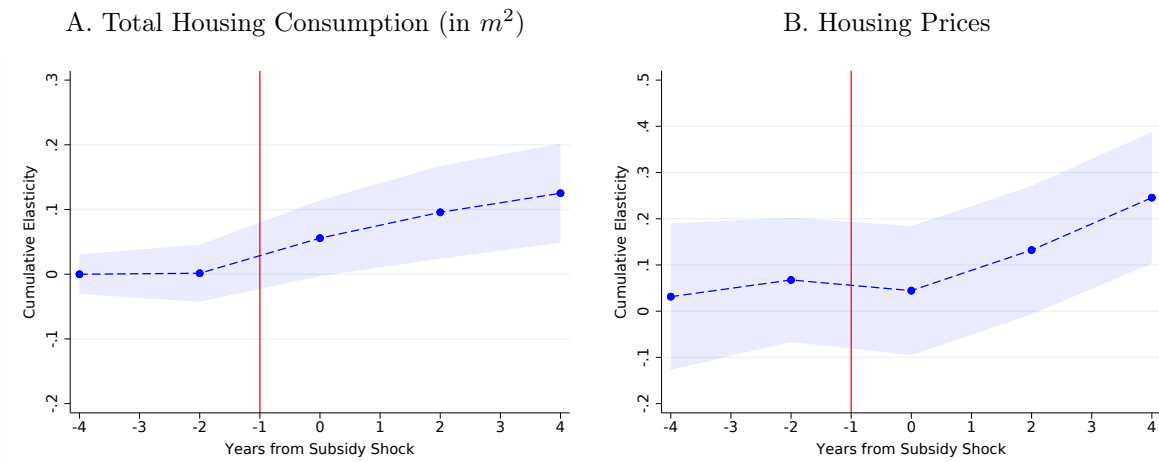
Housing supply and price capitalization Figures 1.8 and 1.9 report the results of our housing supply regressions and show how total m^2 housing consumption and housing prices (in

Figure 1.8: Within-MF Housing Supply and Housing Price Responses



Note: These graphs plot the coefficients β_d^W of Section 1.5 regressions. They show the effect of changes in a municipality's subsidy stock in year 0 on cumulative changes in total m^2 of housing consumed and housing prices starting 6 years prior to the shock year. All changes are relative to MF mean changes. Standard errors are clustered at the MF level. We report the 5% confidence bands.

Figure 1.9: Between-MF Housing Supply and Housing Price Responses



Note: These graphs plot the coefficients β_d^B of Section 1.5 regressions. They show the effect of changes in a MF's subsidy stock in year 0 on cumulative changes in total m^2 of housing consumed and housing prices starting 6 years prior to the shock year. All changes are relative to regional mean changes. Standard errors are clustered at the MF level. We report the 5% confidence bands.

(€/m²) are affected by investment subsidies changes. Panel A of Figure 1.8 reports estimates for housing consumption changes conditional on municipal federation. A 1% increase in the subsidy stock relative to the MF average increase is met with a 0.02% relative increase in total consumed m^2 by 20–65 residents after five years. This estimate is significant at the 1% level. Panel B shows housing price capitalization results: a 1% increase in the subsidy stock is met with an insignificant 0.02% increase in housing prices after five years.

Panel A of Figure 1.9 reports estimates for MF-level mean housing consumption changes conditional on region. A 1% increase in the subsidy stock relative to the regional average increase is met with a 0.13% relative increase in total consumed m^2 by 20–65 residents after five years. This estimate is significant at the 1% level. Panel B again shows housing price capitalization estimate. This time, a 1% increase in the subsidy stock is met with an 0.24% increase in housing prices after five years significant a 1%. Panel B hence comforts our inter-

Table 1.4: Housing Supply Sensitivity Analysis

	(1)	(2)	(3)
	$\Delta \ln \bar{H}_j$	$\Delta \ln \bar{H}_j$	$\Delta \ln \bar{H}_j$
$\Delta \ln \bar{S}_j$	0.027*** (0.007)	0.026*** (0.007)	0.020*** (0.007)
	$\Delta \ln \bar{r}_j$	$\Delta \ln \bar{r}_j$	$\Delta \ln \bar{r}_j$
$\Delta \ln \bar{S}_j$	-0.026 (0.025)	0.015 (0.017)	0.016 (0.017)
Lag $\Delta \ln \bar{S}_j$		Yes	Yes
Baseline Controls			Yes
Observations	99,593	99,593	99,593
	$\Delta \ln \widehat{H}_a$	$\Delta \ln \widehat{H}_a$	$\Delta \ln \widehat{H}_a$
$\Delta \ln \widehat{S}_a$	0.073** (0.035)	0.125*** (0.039)	0.132*** (0.041)
	$\Delta \ln \widehat{r}_a$	$\Delta \ln \widehat{r}_a$	$\Delta \ln \widehat{r}_a$
$\Delta \ln \widehat{S}_a$	0.173** (0.072)	0.246*** (0.073)	0.239*** (0.075)
Lag $\Delta \ln \widehat{S}_a$	Yes	Yes	Yes
Baseline Controls	Yes	Yes	Yes
Time FE X	County	Region	Nation
Observations	5,428	5,428	5,428

Note: This Table reports estimates for $\beta_{d=4}$, i.e., the effect of the shock happening in $d = 0$ on cumulative outcome growth between $d = -6$ and $d = 4$. Lag shocks include all past shocks up to $d = -1$. Controls include baseline 1999 jurisdiction population and density. Standard errors are clustered at the MF level. * significant at 10%; ** significant at 5%; *** significant at 1%.

pretation of subsidy shocks as additional public good amenities that are positively valued by residents. Most importantly, significant housing price capitalization in the between-MF setup and not in the within-MF setup is consistent with the presence of strong within-MF spillovers. It is worth noting that housing price response is an order of magnitude larger than the per capita housing drop in the between-MF case. We argue that housing price responses are more representative of the longer-term adjustment in rental prices facing newcomers as short- to medium-term per capita housing responses might be dampened by adjustment frictions. We use housing prices in our GMM to estimate the model's parameters.

Table 1.4 reports alternative estimates. Our preferred estimates in column (3) correspond to the endpoints of Figures 1.8 and 1.9. They are robust to different specifications.

Overall, these reduced-form results show that households value local public goods as evidenced by migration responses and capitalization in housing prices. They also highlight, in line with our theoretical framework, that the intensity of cross-boundary spillovers influences the magnitude of migration responses to local public good supply shocks.

1.5.3 Robustness Checks

We run a series of robustness checks investigating whether the observed migration patterns could be driven by mechanisms other than responses to changes in positively valued public goods.

A first concern is that changes in local public spending may not be valued by residents in themselves, but may be correlated with changes in housing supply determinants through \mathcal{E}^C . This would be the case if public good shocks were land improvements—new roads, pathways, etc.—of no intrinsic value but destined to welcome social housing units or private housing developments following changes in land use regulation. Residents would migrate towards municipalities experiencing positive housing supply shocks because of lower rents. More generally, if our subsidy shocks are correlated with shocks in the determinants of housing supply it may bias the interpretation of our estimates. As a test to alleviate this concern, we can look back at the rents results of Figure 1.9. A positive housing supply shock would have a negative effect on rents. Our results show the exact opposite suggesting that people indeed value local public goods beyond any correlated shift in the housing supply curve.

A second concern is that migration responses may entirely be driven by the inflow of public

employees necessary to operate the new facilities or services financed by the subsidy shocks. Indeed, our model does not account for public employment. Public goods may be of little value in themselves, but workers may react to public labor demand shocks that would increase wages. It is not conceptually a problem as we could have modelled public good provision as taking public employment as input. The identification of the partial effect of public good supply would still be achieved in the GMM procedure provided that we separately instrument changes in local wages. We nevertheless assess the importance of this channel and show evidence that the public employment effect is marginal. We look at how the share of public employees in the population—crudely measured as public staff payroll divided by total local payroll—evolves around subsidy shocks. Combining this estimate with total population responses and the pre-shock shares of public employees, we conclude that observed behavioral responses coming from public employment only explain approximately 12% of total 20–65 population response (see derivations in appendix A.3).

1.6 Structural Estimation

This section carries out a GMM estimation of the model’s parameters. We derive moments conditions using expressions for changes in residential amenities introduced in Section 1.3. Section 1.5 makes the case for subsidy shocks as a valid instrument in this framework. We provide additional evidence that residential amenity changes are likely mean-independent of subsidy shocks and ground our GMM estimation strategy on these more stringent conditions.

1.6.1 Generalized Method of Moments

As mentioned, we structurally estimate the models’ parameters using a non-linear generalized method of moments. We ground our GMM estimation on moment conditions of the form

$$\mathbb{E} \left[\Delta \ln \mathcal{E}^A \times Z \right] = 0 \tag{1.23}$$

where Z is an instrument uncorrelated with changes in unobserved residential amenities $\Delta \ln \mathcal{E}^A$. Section 1.5 gives support to the following two moment conditions:

$$\begin{aligned} \mathbb{E} \left[\Delta \ln \overline{\mathcal{E}_j^A} \times \overline{Z}_j \right] &= 0 \\ \mathbb{E} \left[\Delta \ln \widehat{\mathcal{E}_a^A} \times \widehat{Z}_a \right] &= 0 \end{aligned} \tag{1.24}$$

where \bar{Z}_j (resp. \widehat{Z}_a) is a variable of residuals obtained from regressing 2007, 2009 and 2010 subsidy shocks $\Delta \ln \bar{S}_j$ (resp. $\Delta \ln \widehat{S}_a$) on a vector of flexible dummies for 1999 baseline characteristics and lagged shocks as in regression (1.21) (resp. (1.22)).

Section 1.5 provides evidence of the absence of pre-trends in the model's endogenous variables, which implies that the empirical counterparts of conditions (1.24) hold in pre-shock periods. To see why, notice that equations (1.19) and (1.20) express changes in residential amenities as a sum of linear functions of observables. Sufficient conditions for (1.24) to hold for pre-shock periods are then $\mathbb{E}[\Delta \ln \bar{Y}_j \times \bar{Z}_j] = 0$ and $\mathbb{E}[\Delta \ln \widehat{Y}_a \times \widehat{Z}_a] = 0$ in pre-shock periods for all observables Y of (1.19) and (1.20), which is precisely what Section 1.5 shows. As is typically done in DiD frameworks, we make the assumption that this absence of correlation between subsidy shocks and amenity changes also holds for post-shock periods, which allows us to estimate model parameters with GMM.

In practice, the GMM procedure looks for the set of parameters that minimizes the empirical counterparts of our moment conditions, keeping the endogenous variables of the model at their observed values in the data. As such, the two moment conditions in (1.24) are not enough on their own to identify all the parameters of our model. Our GMM estimation relies on more moments conditions. In particular, we investigate whether more restrictive conditions of the form $\mathbb{E}[\Delta \ln \mathcal{E}^A | Z] = 0$ hold. It would similarly be supported by evidence that the empirical counterpart of $\mathbb{E}[\Delta \ln Y | Z] = 0$ holds for all outcomes Y in pre-shock periods. The next section provides non-parametric evidence of such relationship between subsidy shock intensity and cumulative outcome changes.

1.6.2 Non-Parametric Evidence

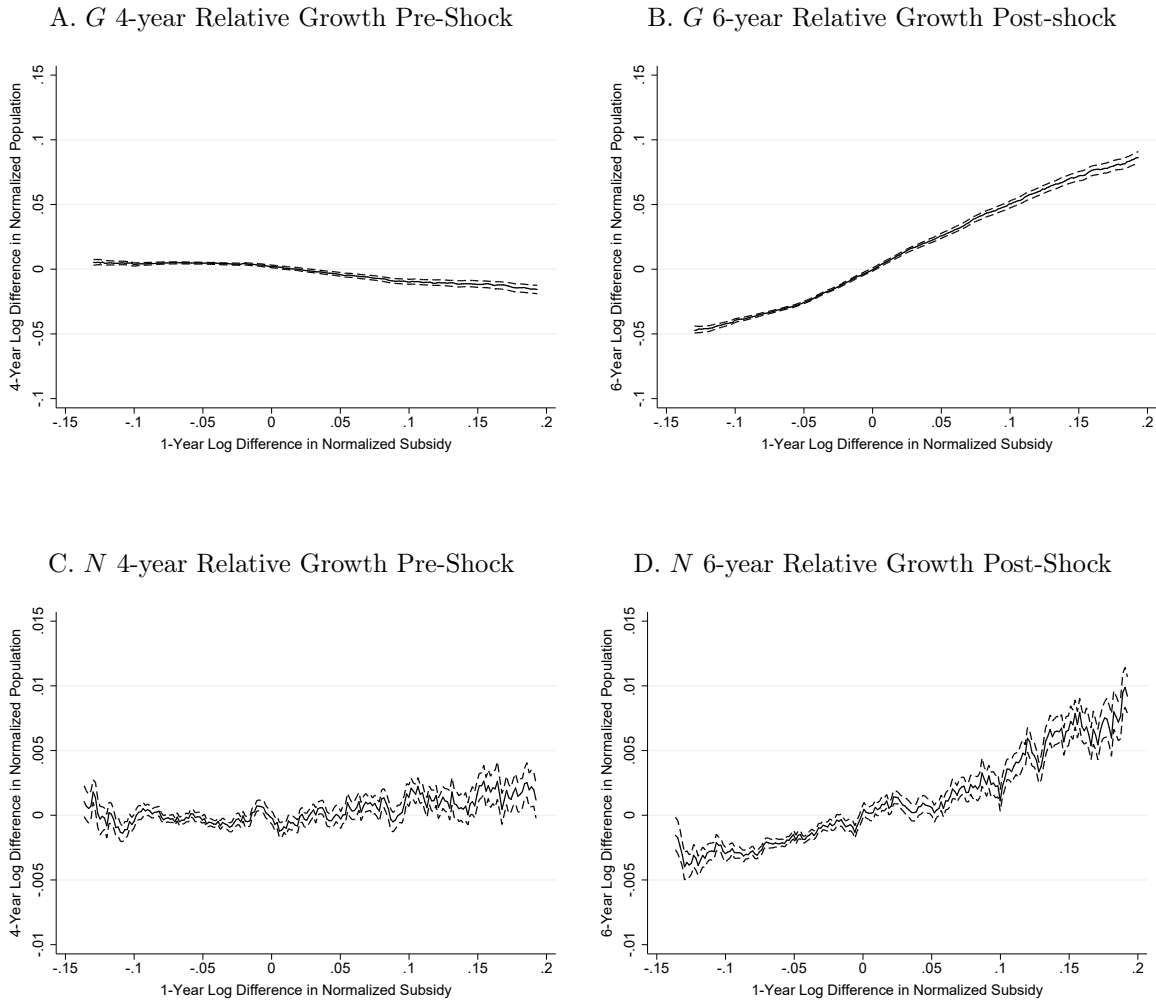
We provide further evidence on the absence of pre-trends by looking at the non-parametric relationship between subsidy shocks on the one hand, and pre- or post-shock cumulative outcome changes on the other hand. We run kernel regressions where the dependent variable is alternatively $\Delta \ln \bar{G}_{j,-6,-2}$, $\Delta \ln \bar{G}_{j,-2,4}$, $\Delta \ln \bar{N}_{j,-6,-2}$ and $\Delta \ln \bar{N}_{j,-2,4}$. The explanatory variable is $\bar{Z}_{j,-1,0}$, i.e., the variable of residuals obtained from regressing 2007, 2009 and 2010 shocks $\Delta \ln \bar{S}_{j,-1,0}$ on the set of baseline controls and lagged shocks.²³ Figure 1.10 presents

²³The fitted values and confidence bands are computed from running kernel regressions of the dependent variable on these initial residuals and on 1000 additional samples of residuals. We generate synthetic residuals using the wild cluster bootstrap procedure proposed in Cameron et al. (2008). We assume that errors are correlated within clusters which we take to be counties. Each cluster randomly draws a +1/−1 coefficient with probability 0.5 and all residuals of a same cluster are multiplied by the same coefficient. These synthetic

the results.

The dependent variable in Panel A (resp. in Panel C) is $\Delta \ln \bar{G}_{j,-6,-2}$ (resp. $\Delta \ln \bar{N}_{j,-6,-2}$). Panels A and C show again that the average relationship between subsidy shock and pre-shock outcome growth is close to zero. In addition, they offer evidence that municipality expected

Figure 1.10: Kernel Regressions: Within-MF Changes



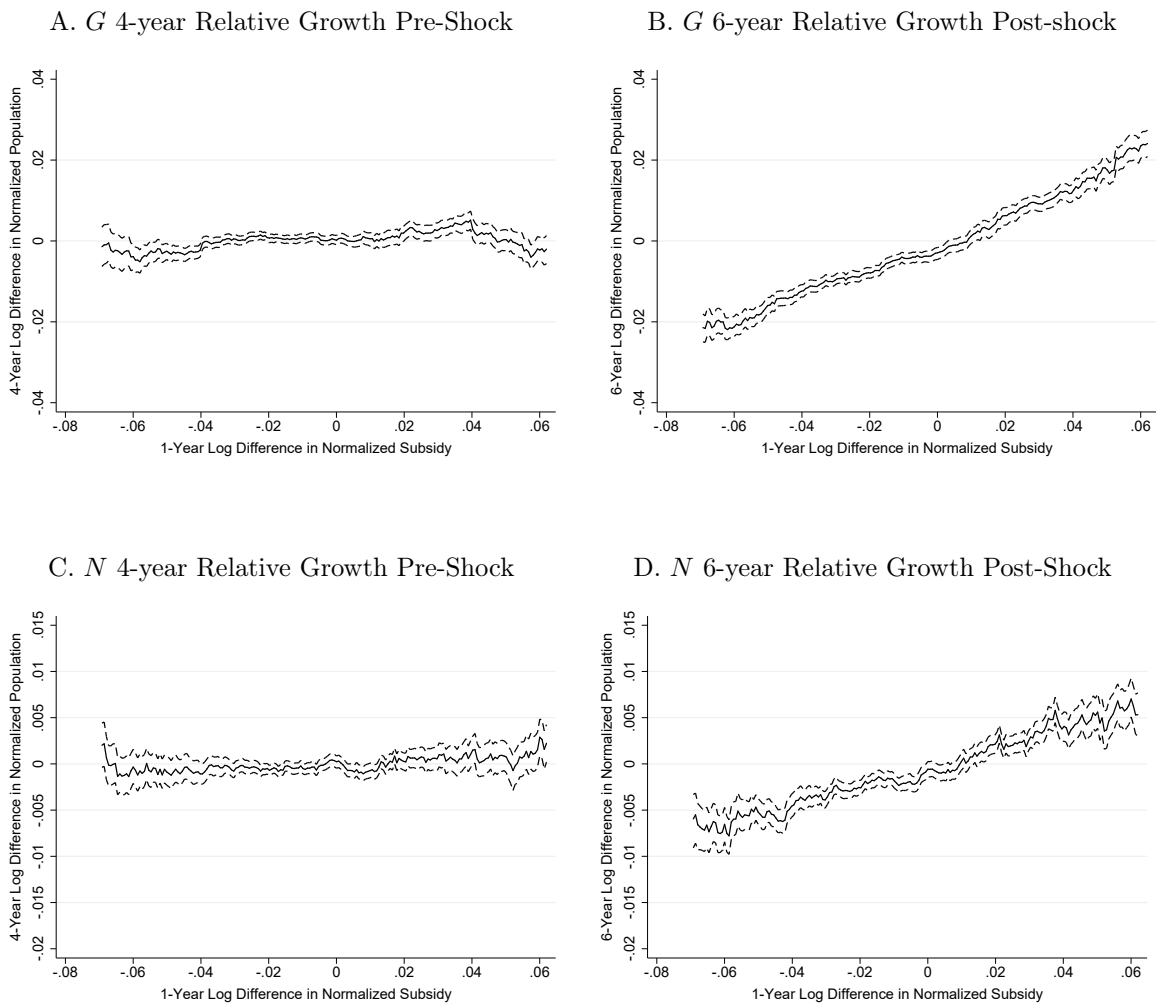
Note: This Figure shows the outcomes of non-parametric regressions where the explanatory variable is the within-MF relative subsidy shock. Dependent and explanatory variables are first residualized with respect to bins of fixed 1999 characteristics and lagged shocks. Bias-corrected fitted values and 1% confidence band are computed based on 1,000 bootstrap replications.

residuals are then added back to the original fitted values. We run the first-step regression again on this pseudo-sample and store the coefficient estimates. We do this 1000 times. We then generate 1000 pseudo-samples of residuals by fitting each saved model on the original data and saving the residuals. The graphs report the pivotal bootstrap confidence band and the bias-corrected fitted values.

outcomes grow at the same rate as their MF geometric average *conditional on shock intensity*. In mathematical terms, this translates into $\mathbb{E} [\Delta \ln \bar{Y}_{j,-6,-2} | \bar{Z}_{j,-1,0}] = 0$ for $Y \in \{G, N\}$. This mean-independence property will be central to our GMM analysis.

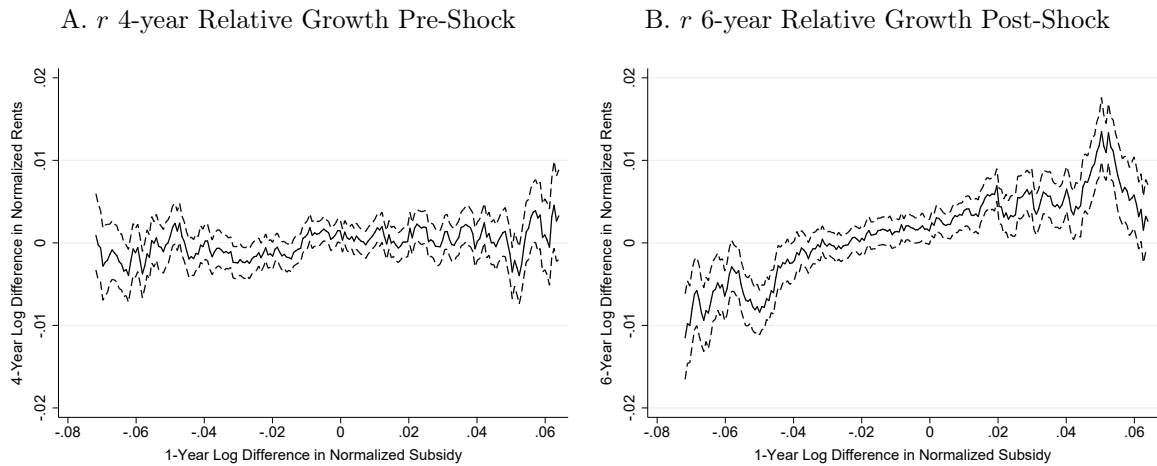
In Panels B and D, the dependent variables are $\Delta \ln \bar{G}_{j,-2,4}$ and $\Delta \ln \bar{N}_{j,-2,4}$. They offer reassurance that the average effects reported in the Figures of Section 1.5 are not driven by outliers. Public good and migration responses are visible across the full distribution of subsidy shocks.

Figure 1.11: Kernel Regressions: Between-MF Changes (1/2)



Note: This Figure shows the outcomes of non-parametric regressions where the explanatory variable is the between-MF relative subsidy shock. All changes are relative to mean regional changes. Dependent and explanatory variables are first residualized with respect to bins of fixed 1999 characteristics and lagged shocks. Bias-corrected fitted values and 1% confidence band are computed based on 1,000 bootstrap replications.

Figure 1.12: Kernel Regressions: Between-MF Changes (2/2)



Note: This Figure shows the outcomes of non-parametric regressions where the explanatory variable is the between-MF relative subsidy shock. All changes are relative to mean regional changes. Dependent and explanatory variables are first residualized with respect to bins of fixed 1999 characteristics and lagged shocks. Bias-corrected fitted values and 1% confidence band are computed based on 1,000 bootstrap replications.

Figure 1.11 provides similar evidence on how $\Delta \ln \widehat{G}_a$ and $\Delta \ln \widehat{N}_a$ change with $\widehat{Z}_{a,-1,0}$. Panels A and C show the absence of correlation between subsidy shock and pre-shock normalized outcome growth which similarly translates into $\mathbb{E} \left[\Delta \ln \widehat{Y}_{a,-6,-2} | \widehat{Z}_{a,-1,0} \right] = 0$ for $Y \in \{G, N\}$. Panels B and D offer additional evidence that responses to subsidy shocks are observed all along the distribution of shock intensity. Finally, Figure 1.12 similarly shows that housing prices $\Delta \ln \widehat{r}_a$ evolve in a similar fashion in different cells of shock intensity in pre-shock periods and that the price response is coming from the full range of shock values.

These graphs also show that changes in jurisdictions' populations and rents are not necessarily proportional to changes in local public goods. Our model supports these non-proportional responses. For example, jurisdictions' proportional responses to changes in local public goods may differ because of heterogeneity in housing supply elasticity or production technology for the traded good. In addition, jurisdictions' budget constraints introduce non-linearity in endogenous variables' responses to subsidy shocks. Importantly, these non-proportional responses are precisely what enables us to add more moment conditions to (1.24).

1.6.3 Moment Conditions

The non-parametric evidence in Section 1.6.2 showed that subsidy shocks likely satisfy mean-independence conditions of the form $\mathbb{E}[\Delta \ln \mathcal{E} | Z] = 0$. These are more restrictive than the usual conditions of zero correlation between instrument and unobserved fundamentals $\mathbb{E}[\Delta \ln \mathcal{E} \times Z] = 0$. Indeed, with mean-independence of the unobserved fundamentals with respect to our original instrument Z , any function of Z may be used as an additional instrument (see Wooldridge 2010). We apply this property to a discrete number of indicator functions as in Ahlfeldt et al. (2015): we partition the empirical distribution of \bar{Z} (or \widehat{Z}) into subintervals of equal range. We define \mathbb{I}_m^W (resp. \mathbb{I}_m^B) the indicator function equal to one if \bar{Z} (resp. \widehat{Z}) belongs to the subinterval m of the partition of \bar{Z} (resp. of \widehat{Z}).

We can now state all our moment conditions:

$$\begin{aligned} \mathbb{E} \left[\Delta \ln \overline{\mathcal{E}}_j^A \times \mathbb{I}_m^W \right] &= 0 \quad \text{for all } m \in \{1, \dots, M\} \\ \mathbb{E} \left[\Delta \ln \overline{\mathcal{E}}_j^A \times \bar{Z}_j \right] &= 0 \end{aligned} \tag{1.25}$$

using the within-MF expression (1.19) for residential amenities and

$$\begin{aligned} \mathbb{E} \left[\Delta \ln \widehat{\mathcal{E}}_a^A \times \mathbb{I}_m^B \right] &= 0 \quad \text{for all } m \in \{1, \dots, M\} \\ \mathbb{E} \left[\Delta \ln \widehat{\mathcal{E}}_a^A \times \widehat{Z}_a \right] &= 0 \end{aligned} \tag{1.26}$$

using the between-MF expression (1.20) for residential amenities. These additional moments conditions provided by indicator variables $\{\mathbb{I}_m^W, \mathbb{I}_m^B\}$ bring additional information for the GMM estimation only if they are not collinear. Section 1.6.2 provides such evidence and discusses the sources of non-linearity driving the endogenous variables' non-proportional responses.

All outcome changes plugged in (1.25) and (1.26) are the eight-year relative cumulative growths between year -4 and year 4 around the 2007, 2009 and 2010 subsidy shocks. The identifying assumption is that subsidy shocks and indicator variables $\{\mathbb{I}_m^W, \mathbb{I}_m^B\}$ are uncorrelated with jurisdiction or MF changes in unobserved residential amenities. As mentioned, DiD results of Section 1.5.2 and their decomposition in Section 1.6.2 suggest that this is indeed the case.

We simultaneously estimate the model's parameters using a two-step non-linear GMM procedure. The GMM estimation looks for the set of parameters that minimizes the empirical counterpart of moment conditions (1.25) and (1.26), keeping the endogenous variables of the model at their observed values in the data. As in Ahlfeldt et al. (2015), our moment conditions

may not uniquely identify the model's parameters as the objective function might not be globally concave. In practice, the GMM objective function is well behaved and we find a global minimum.

To make sure that the arbitrarily chosen number of subintervals for the partition of \bar{Z} and \hat{Z} does not affect our results, we try alternative specifications with different numbers of moments. However, the estimation requires a sufficient number of subintervals to pick up enough variation in jurisdiction responses to subsidy shocks. We report parameter estimates for specifications varying between 40 and 60 subintervals for each calibrated value of the housing consumption share.

1.6.4 Estimation Results

We report estimates of our structural parameters for different calibrated values of the housing consumption share α . Calibrating the consumption share allows us to improve the robustness of our estimates given the multiplicative fashion in which it interacts with other parameters. Tables 1.5, 1.6, 1.7 and 1.8 report estimates of the model's parameters for $\alpha \in \{0.25, 0.30, 0.35, 0.40\}$.

We estimate a taste parameter for public goods ϕ that varies between 0.11 and 0.25 according to the calibrated housing share. For a given value of α , our estimate is robust for different number of moments. Given the Cobb-Douglas specification of preferences, the interpretation is that municipalities spend between 11% and 25% of their total resources on public goods. This estimate is typically increasing with the calibrated α . Indeed, a higher taste for public goods is needed to compensate for given housing price changes when the taste for housing is higher, holding constant migration and public good changes. Parameter ϕ is the only parameter which is directly comparable with estimates from the existing literature, which we report in Table 1.9. Our estimates fall in the range of existing estimates.

Estimates for inverse household mobility σ lie between 0.03 and 0.14. These estimates are much lower than those found for instance in [Serrato and Wingender \(2011\)](#) or [Diamond \(2016\)](#). First, notice that the size of the considered French jurisdictions is much smaller than the geographic unit of these studies, typically the MSA. Mobility between locations is then expected to be different. Second, it is interesting to note that given modeling assumptions, parameter σ is isomorphic to any combination of parameters $\sigma - \tilde{\kappa}$ where $\tilde{\kappa}$ would capture positive agglomeration externalities from increased density beyond those transiting through

increased public goods and potential changes in wages. Hence, our estimated σ is actually capturing mobility frictions net of all non-public and non-productive agglomeration effects, e.g., endogenous residential amenities such as in [Diamond \(2016\)](#).

We find substantial public good spillovers between jurisdictions of a same municipal federation. Estimates for δ all lie between 0 and 0.08. While they are typically decreasing with the calibrated α , the relationship is not trivial. When taste for housing increases, lower benefit spillovers are needed to explain observed housing price increases, i.e., location must matter for the enjoyment of public goods. However, higher congestion spillovers can also explain observed price changes, because additional congestion brought by new residents matters less than congestion caused by residents of neighbor jurisdictions. Our estimates show that the second effect prevails.

Recall that $\delta = 0$ means full spillovers within a municipal federation, while $\delta = 1$ means no spillovers. In line with our previous empirical evidence, our estimates suggest that spillovers are very strong in the examined French setting. A back-of-the-envelope calculation for the average municipal federation size—27 member municipalities—implies that local public spending in a municipality accounts for 4–11% of the bundle of local public goods enjoyed by its residents, and that local public spending in its average neighbor municipality—belonging to the same municipal federation—accounts for 3.2–3.5% of this bundle.²⁴ These large estimates are unsurprising given the large number of French jurisdictions and the fact that the only restriction to benefiting from local public goods in other municipalities is travel distance—there is almost no legal restriction such as zoning.

Our estimates for public good congestion are the least robust across specifications. However, we find significant estimates for $\kappa\phi$ that are typically below estimates for ϕ . This points to local public services being not fully rival and subject to economies of scale, and to the existence of other sources of local inefficiencies due to fiscal agglomeration effects. Indeed, these estimates suggest that the cost of providing public services increases less than one for one with population, making denser jurisdictions more attractive from a public amenity perspective.

²⁴The weight put on own public goods is $\delta + \frac{1}{|a|}(1 - \delta)$ where $|a|$ is the number of member jurisdictions in the MF. We take the mean number of member jurisdictions, i.e., 27 for this computation.

Table 1.5: Structural Parameters for $\alpha = 0.25$

# Subintervals	40	42	44	46	48	50	52	54	56	58	60
Spillovers (δ)	0.077	0.051	0.071	0.024	0.029	0.034	0.063	0.056	0.031	0.063	0.042
	0.011	0.005	0.008	0.009	0.012	0.014	0.007	0.007	0.017	0.007	0.009
PG Taste (ϕ)	0.114	0.150	0.139	0.142	0.147	0.138	0.155	0.175	0.153	0.163	0.148
	0.007	0.005	0.005	0.004	0.009	0.012	0.006	0.007	0.009	0.012	0.007
Mobility (σ)	0.113	0.103	0.142	0.069	0.065	0.079	0.125	0.124	0.080	0.118	0.099
	0.009	0.011	0.006	0.008	0.014	0.012	0.007	0.005	0.013	0.006	0.004
PG Congestion ($\kappa\phi$)	0.056	0.113	0.065	0.134	0.166	0.118	0.094	0.131	0.174	0.144	0.061
	0.013	0.021	0.012	0.012	0.022	0.027	0.017	0.010	0.024	0.016	0.038

Note: This Table reports our structural parameters estimation for $\alpha = 0.25$. Standard errors are computed using a wild cluster bootstrap procedure based on 1,000 replications.

Table 1.6: Structural Parameters for $\alpha = 0.30$

# Subintervals	40	42	44	46	48	50	52	54	56	58	60
Spillovers (δ)	0.073	0.037	0.053	0.016	0.028	0.031	0.056	0.041	0.015	0.052	0.035
	0.009	0.005	0.004	0.007	0.013	0.009	0.008	0.004	0.014	0.008	0.006
PG Taste (ϕ)	0.132	0.178	0.162	0.165	0.165	0.154	0.173	0.194	0.183	0.190	0.168
	0.004	0.009	0.009	0.005	0.007	0.011	0.009	0.007	0.012	0.012	0.007
Mobility (σ)	0.104	0.083	0.112	0.048	0.046	0.069	0.104	0.098	0.067	0.105	0.080
	0.007	0.013	0.006	0.008	0.013	0.009	0.009	0.007	0.016	0.010	0.006
PG Congestion ($\kappa\phi$)	0.063	0.152	0.100	0.160	0.168	0.138	0.109	0.166	0.197	0.161	0.064
	0.010	0.023	0.013	0.018	0.023	0.025	0.013	0.011	0.021	0.024	0.024

Note: This Table reports our structural parameters estimation for $\alpha = 0.30$. Standard errors are computed using a wild cluster bootstrap procedure based on 1,000 replications.

Table 1.7: Structural Parameters for $\alpha = 0.35$

# Subintervals	40	42	44	46	48	50	52	54	56	58	60
Spillovers (δ)	0.047	0.027	0.043	0.000	0.000	0.025	0.037	0.040	0.012	0.025	0.015
	0.008	0.005	0.005	0.008	0.007	0.006	0.006	0.010	0.007	0.006	0.010
PG Taste (ϕ)	0.155	0.188	0.179	0.200	0.196	0.177	0.194	0.211	0.193	0.219	0.188
	0.009	0.010	0.005	0.007	0.005	0.015	0.006	0.013	0.010	0.008	0.007
Mobility (σ)	0.092	0.071	0.088	0.031	0.014	0.063	0.092	0.086	0.045	0.071	0.065
	0.010	0.010	0.006	0.011	0.013	0.007	0.006	0.011	0.008	0.008	0.005
PG Congestion ($\kappa\phi$)	0.092	0.156	0.123	0.200	0.225	0.176	0.154	0.179	0.229	0.187	0.142
	0.023	0.020	0.017	0.013	0.009	0.043	0.019	0.019	0.015	0.012	0.033

Note: This Table reports our structural parameters estimation for $\alpha = 0.35$. Standard errors are computed using a wild cluster bootstrap procedure based on 1,000 replications.

Table 1.8: Structural Parameters for $\alpha = 0.40$

# Subintervals	40	42	44	46	48	50	52	54	56	58	60
Spillovers (δ)	0.035	0.019	0.037	0.007	-0.001	0.014	0.025	0.027	0.017	0.023	0.017
	0.005	0.005	0.007	0.004	0.004	0.008	0.003	0.009	0.008	0.007	0.004
PG Taste (ϕ)	0.181	0.221	0.183	0.214	0.221	0.205	0.230	0.254	0.232	0.252	0.228
	0.009	0.008	0.009	0.006	0.006	0.021	0.009	0.010	0.012	0.012	0.007
Mobility (σ)	0.070	0.050	0.074	0.033	-0.003	0.046	0.078	0.063	0.042	0.063	0.051
	0.007	0.014	0.005	0.007	0.012	0.015	0.005	0.008	0.015	0.011	0.010
PG Congestion ($\kappa\phi$)	0.126	0.217	0.140	0.230	0.252	0.156	0.193	0.226	0.262	0.171	0.163
	0.017	0.030	0.022	0.011	0.027	0.041	0.016	0.013	0.017	0.027	0.033

Note: This Table reports our structural parameters estimation for $\alpha = 0.40$. Standard errors are computed using a wild cluster bootstrap procedure based on 1,000 replications.

Table 1.9: Structural Parameters Calibration and Estimation

Parameter	Values from literature	Treatment in this paper
α	Diamond (2016) ^a : 43% for US non-college workers and 46% for US college workers	Calibration from literature
	Combes et al. (2018) ^e : 0.314 for owners and 0.352 for renters for France	
	Paris Lyon, Lille, Marseille Pop > 200,000 Pop ≤ 200,000	
	owners 0.344 0.344 0.304 0.293	
renters 0.369 0.367 0.382 0.285		
ϕ	Serrato and Wingender (2011) ^b : [0.391,0.502] for US non-college workers and [0.228,0.267] for US college workers	Estimation with GMM
	Fajgelbaum et al. (2015): [0.11,0.23] for the US	
	Serrato and Wingender (2016) ^d : 0.26 for the US	
	Diamond (2016) ^a : 0.03 for US non-college workers and 0.32 for US college workers	
σ	Serrato and Wingender (2011) ^b : [0.342, 0.399] for US non-college workers and [0.350, 0.376] for US college workers	Estimation with GMM
	Diamond (2016) ^a : 0.24 for US non-college workers and 0.47 for US college workers	
η	Diamond (2016) ^a : 0.21 with a standard deviation of 0.22 for the US	Calibration from literature
	Serrato and Wingender (2011) ^b : [0.407,0.813] for the US	
	Combes et al. (2018) ^e : 0.208 with most alternative estimates being between 0.15 and 0.30 for France	
ζ		Calibration with our data
φ		Calibration with our data
κ		Estimation with GMM
δ		Estimation with GMM

^a α : the author uses US CEX survey data. ϕ : $1.012/2.116 = \frac{\phi}{1-\phi}$ for unskilled workers and $0.274/4.026 = \frac{\phi}{1-\phi}$ for skilled workers. ϕ is a structural parameter capturing workers' valuation of all amenities compared to the traded good and is not public good specific. σ : we take σ to be the inverse of the structural wage coefficient in the author's favourite specification (3) which gives $1/0.4026 = 0.24$ for US non-college workers and $1/2.116 = 0.47$ for US college workers. η : inverse housing supply elasticity of 0.21 with standard deviation of 0.22.

^b $\eta = 0.813$ when using housing prices and 0.407 when using rents. Author's preferred specification is non-linear.

^dFrom Fajgelbaum et al. (2015) literature review on structural parameters. See their Table A.17.

^eUse family expenditure survey for α values.

1.7 Welfare Implications

Our empirical analysis in Sections 1.5 and 1.6 documents strong public good spillovers across jurisdiction borders. Decentralized public good provision is likely to be sub-optimally low, i.e., there may be welfare gains from coordination. While current subsidies from counties, regions and the central government may carry some Pigovian flavour, it is unlikely that they fully tackle spillover inefficiency. In this section, we focus on a simplified version of our model and simulate the welfare impact of an administrative reform merging together all municipalities belonging to a same municipal federation. As mentioned in Section 1.2, these groupings were historically introduced precisely to deal with spillovers and economies of scale. However, member jurisdictions still maintained a high degree of independence and it is a natural check to investigate whether this residual autonomy is at the root of significant deadweight loss.

We first back out the fundamentals of the model $\{\mathcal{E}_j^A, \mathcal{E}_j^C\}$ from the model's equilibrium variables that we observe as econometricians in the year 2014, and from parameters that we have estimated or calibrated from the literature. We then make several assumptions to ensure that the equilibrium in our simplified setup is tractable. Motivated by the suppression of the French local business tax in 2011, we first assume that local public revenues only come from the residence tax. In this simplified version, we also assume away public subsidies from the central government. We make a series of other assumptions: (i) residents are myopic and do not anticipate migration responses when they vote for public goods and taxes; (ii) the public good production function is the identity, i.e., $\Gamma(x) = x$; (iii) housing supply elasticities η_j are constant across jurisdictions; (iv) local productivity θ_j^Y does not depend on local public goods and (v) the geographic fundamentals that we back out from the observed 2014 allocation still define the geographic landscape in this simplified model. This version of the model has a unique and tractable equilibrium given parameters $\{\alpha, \kappa, \sigma, \phi, \delta, \eta\}$ and geographic fundamentals $\{\mathcal{E}_j^A, \mathcal{E}_j^C\}$. While this setting does not allow us to evaluate a reform that applies to the observed situation, it helps illustrate the magnitude of potential welfare gains.

Prior to the reform, residents vote for public goods as in Section 1.3.6 absent subsidies and business taxation. They choose the housing tax rate τ_j^h and the level of public good G_j that maximize v_j under the budget constraint $r_j N_j h_j \tau_j^h = G_j$. Following the merger, residents vote for a new level of public spending pooling tax bases of all former member jurisdictions. To ensure that welfare remains comparable between pre- and post-reform situations, we assume that pre-reform jurisdictions are still relevant for labor and housing markets. The only change

is the degree of cooperation in the provision of local public goods, now chosen at the MF level. Residents of $j \in a$ hence choose the housing tax τ_a^h and the level of public good G_a that maximize v_j under the budget constraint $G_a = \sum_{j' \in a} r_{j'} N_{j'} h_{j'} \tau_a^h$.²⁵ We assume that the total amount of local public goods is then allocated to all jurisdictions of a in proportion to their pre-reform share of the total quantity of public good in the municipal federation. This allocation rule allows us to center our welfare analysis on efficiency gains and neutralizes the redistributive channel that would be present with, say, an egalitarian allocation. Our analysis would also work with alternative allocation rules but the welfare channels would be more intricate.

We note X^o the variables observed in 2014, which we assume are the equilibrium variables of the full model. In the simplified model, we call X^n the pre-reform equilibrium variables and X^m the post-reform equilibrium variables. Local public good supply before the reform is characterized by

$$1 + \tau_j^n = 1 + \frac{\phi(\delta + \frac{1-\delta}{|a_j|})}{(1-\phi)\alpha} \quad (1.27)$$

$$G_j^n [1 + \tau_j^n] = \Phi \tau_j^n N_j^n w_j^n \alpha$$

where Φ is a constant defined in Appendix A.4. After the merger, local public good supply is characterized by

$$1 + \tau_a^m = 1 + \frac{\phi}{(1-\phi)\alpha}$$

$$G_j^m = \Phi \frac{G_j^n}{\sum_{j' \in a_j} G_{j'}^n} \sum_{j' \in a_j} \frac{\tau_{j'}^m}{1 + \tau_{j'}^m} N_{j'}^m w_{j'} \alpha \quad (1.28)$$

Using backed out fundamentals, we can express all endogenous variables in these two simulated cases as functions of 2014 observables. We assume that social welfare is²⁶

$$W = \mathbb{E} \left[\max_j \ln U(\mathcal{C}, \mathcal{G}, i, j) \right] \quad (1.29)$$

Because idiosyncratic preferences are distributed Extreme Value Type-I, the welfare change

²⁵The result of this optimization problem is the same for any jurisdiction j in a , see Appendix A.4.

²⁶The choice of a utilitarian welfare criterion is akin to considering that policy choices are made behind the “veil of ignorance”. Increasing welfare is then equivalent to increasing the expected utility of ex-ante homogeneous agents. Evaluating the proposed reform through the lens of an ex-post Pareto criterion will likely change its desirability.

associated with the reform is

$$\begin{aligned}\Delta W &= W^m - W^n \\ &= \sigma \ln \left(\frac{\sum_j \exp(v_j^m/\sigma)}{\sum_j \exp(v_j^n/\sigma)} \right)\end{aligned}\tag{1.30}$$

In the model, utility is homogeneous to euros so we interpret these welfare changes as percentage changes of a money metric. All detailed equations are given in Appendix A.4.

In our baseline scenario, we simulate the welfare change taking $\alpha = 0.3$ for the housing consumption share and $\eta = 0.2$ for the housing supply elasticity. These two values correspond to the literature’s central estimates for France (see Table 1.9). For each parameter of $\{\kappa, \sigma, \phi, \delta\}$, we use its average estimation across specifications of Section 1.6. Because this simulation is based on estimated parameters and fundamentals, we report the associated Monte Carlo standard errors that we compute using the parameters’ estimated variance-covariance structure.

Table 1.10: Welfare Change Estimates

(a) Welfare Change for $\alpha = 0.30$			(b) Welfare Change for $\eta = 0.20$		
η	ΔW mean	ΔW sd	α	ΔW mean	ΔW sd
0.15	60%	19%	0.25	40%	13%
0.20	63%	21%	0.30	63%	21%
0.25	66%	23%	0.35	100%	29%
			0.40	137%	35%

Note: These Tables report the welfare impact of a merger of all French jurisdictions at the municipal federation level in our simplified model, holding constant either α or η . Standard errors are computed based on 1,000 Monte Carlo simulations.

In our central specification, we estimate that a reform that would fully merge jurisdictions at the existing federation level would increase welfare by around 60%. This welfare gain is significant at the 1% level. Tables 1.10a and 1.10b report welfare change estimates for alternative values of η and α . They are largely robust across values of η . However they vary between 40% and 137% when we plug alternative housing consumption shares holding $\eta = 0.20$. This is expected as our GMM estimates for spillovers and taste for public goods are increasing with the calibrated α , which makes welfare gains from a coordination reform increasing with α .

This strong result is unsurprising given our spillover estimates, but should be interpreted as an upper bound. First, it is obtained in a hypothetical context without public subsidies. Because current subsidies may already be partly correcting spillover externalities, we expect that the welfare gains from a reform that would apply to the observed situation holding constant observed subsidies would be smaller. We also emphasize that by abstracting from the political deadweight loss that may arise under a more centralized regime, our analysis only investigates one side of the centralization efficiency trade-off. More work is needed to finely assess how much the inability to tailor policies to local needs as well as other potential political frictions would decrease overall welfare gains.

1.8 Conclusion

This paper develops a simple yet flexible framework to test for potential welfare gains from centralization of public goods provision. We first build a spatial equilibrium model with endogenous public goods causing consumption spillovers across jurisdictions' borders. The binary structure for spillover spatial decay makes it easily amenable to empirical analysis. Our model shows how one can exploit differential behavioral responses to shocks in local public goods at different geographic levels to uncover the intensity of public good spillovers.

We then bring new insights on spillovers in the fragmented French institutional context by providing reduced-form evidence of migration and housing price responses to changes in local public goods. We estimate our model with GMM and find substantial public good spillovers, corroborating our reduced-form evidence. A jurisdiction's public goods account for approximately 4–11% of the local public good bundle enjoyed by its residents, and public spending in each neighbor municipality — belonging to the same municipal federation — approximately accounts for an average 3.2–3.5% of this bundle.

In a final exercise, we simulate the effect of a reform redefining jurisdictions' administrative boundaries in a simpler version of our model. Although we do not estimate the cost of centralization, our results suggest that increased coordination in the provision of public goods may substantially improve welfare.

Chapter 2

Optimal Spatial Policies with Public Goods and Unobserved Location Preferences*

We study the normative implications of the interplay between public good agglomeration economies and heterogeneous location preferences in an economic geography framework. Using data on French cities, we first present new stylized facts suggesting scale economies in the consumption of local public goods. We then build a spatial equilibrium model with mobile workers and endogenous public goods and characterize the transfers implementing efficient population distributions. We show that heterogeneity in location preferences increases the equity cost of migration — spatial transfers — relatively to its efficiency benefits — agglomeration gains — such that Pareto-improving reforms may not always exist. Investigating the Pareto efficiency of current French transfers, we empirically show that ignoring location preferences leads to recommending reforms that mistakenly redistribute towards the densest and richest places. Finally, we argue that the interaction between public good scale economies and location preferences may raise horizontal equity concerns about public good consumption, and find evidence that the French planner compensates low-density areas beyond a utilitarian objective.

*This chapter is based on a joint work with Benjamin Carantino.

“*Vivre et travailler au Pays*”

(“Living and working locally”)

— *Bonnets Rouges* (Brittany, 2013)

2.1 Introduction

The economic geography literature has given location preferences various interpretations, from capturing mobility frictions to representing genuine geographic tastes. While mobility costs have historically decreased (see [Combes and Lafourcade 2005](#)) and may have become negligible over one’s lifetime in developed countries, genuine preferences capture richer attachments to locations — e.g., to birthplace, social or natural amenities — and need not vanish in the long run. Indeed, as highlighted by [Lévy et al. \(2018\)](#), unprecedented spatial liberty in contemporary high-mobility societies likely led to the prevalence of such idiosyncratic criteria in residential choices.¹

From a public finance perspective, this freedom of movement raises concerns about the efficiency of population distribution due to potential economies of scale in the consumption of local public goods — the public good agglomeration effect. Indeed, because migrating households do not internalize their positive impact on cities’ tax revenues nor the extra crowding they bring to existing public goods, migration comes with both a positive fiscal externality and a negative congestion externality. If these public goods are partly funded through taxes on local profits accruing to landowners nationally, taxation may also be inefficient. If these externalities do not offset each other, there may be room for Pareto-improving government transfers that raise everyone’s welfare by reaching more efficient population distribution and public good provision.

However, the interplay of these public good externalities with households’ geographic preferences likely determines whether spatial policies can actually achieve efficiency gains. Because the government is unable to observe individual location preferences, transfers are bound to be place-based. Spatial transfers then create direct winners (infra-marginal residents of subsidized places) and losers (infra-marginal residents of taxed places). They symmetrically create indirect losers and winners through net agglomeration gains and changes in aggregate profits. One may fear that the overall welfare effect of spatial transfers always be negative, in partic-

¹[Fauchille \(2016\)](#) and [Lévy and Fauchille \(2017\)](#) ran surveys in France and Switzerland showing that a very large majority of households mention their residence as an idiosyncratic choice among a wide range of available options, from city centers to low-density peripheries.

ular for those taxed residents that are strongly attached to their city. It is then natural to wonder under what circumstances Pareto-improving policies do exist.

Geographic preferences also raise questions about the fairness of the spatial equilibrium. Spatial transfers typically achieve social justice goals, usually by redistributing towards low-productivity areas.² We argue that public good agglomeration economies—in addition to justifying intervention on efficiency grounds—can provide further motivation for redistributive spatial transfers in the presence of location preferences. Individuals that are strongly attached to places destined to be low-density consume less public goods than residents of high-density places when net agglomeration externalities are positive, sometimes for similar individual contributions to local budgets. This mechanism may help explain revived concerns for “spatial justice” and feelings of tax inequity in low-density areas, a recent expression of which may be France’s 2018 Yellow Vests (*Gilets Jaunes*) movement.^{3,4} If society values some form of horizontal equity—e.g., workers with similar incomes or contributions should get similar public good benefits—means-tested taxes and transfers will in general not be sufficient to carry out redistribution: heterogeneity in population density conditional on income will require transfers that are expressly place-based to redistribute from high- to low-density places. Whether or not idiosyncratic geographic attachments and their welfare consequences are to be compensated for is obviously a highly sensitive question.

In this paper, we study the welfare implications of the interplay between public good agglomeration economies and heterogeneous location preferences. We first present new stylized facts on scale economies in the consumption of local public goods by exploiting a unique combination of longitudinal administrative datasets on French cities with detailed information on municipal financial accounts, detailed tax revenues, local income data, municipal population and land use data over the period 2002–2014. Importantly and unlike existing studies, our data covers public assets in addition to yearly public spending. Total per capita public spending decreases with population with an elasticity of 0.32 in our main specification. Although these patterns are better interpreted as associations in the data rather than causal effects, they

²Means-tested taxes and transfers can be viewed as implementing geographic equity objectives when location determines the wage (see [Albouy 2009](#)).

³Using data on 2,900 French households, [Spire and Bernard \(2019\)](#) document that, other things being equal, the probability to view the tax system as unfair is 35% to 68% larger for inhabitants of small cities (2,000 to 20,000 inhabitants) than for inhabitants of large cities (200,000 inhabitants and more) and 40% larger in rural areas overall.

⁴In low-density areas, claims for a “right to stay” have also been documented, e.g., by [Spire \(2018\)](#) who noted the above-quoted slogan during Brittany’s Red Hats (*Bonnets Rouges*) 2013 regional social unrest.

suggest substantial gains from density through the public good consumption amenity.

We then develop an economic geography model with endogenous local public goods where mobile workers have heterogeneous preferences for the various locations. As we wish to emphasize the spatial nature of efficiency and equity concerns, we assume away differences in other kinds of preferences and in skills, endowments or ownership. Workers consume a freely traded private good and a local non-traded private good such as housing. Local jurisdictions provide public goods according to the outcome of a unanimous vote from residents. All three goods are supplied by locally competitive sectors. To have our model match the French institutional setting, we assume that jurisdictions fund public goods through lump-sum taxes on resident households and property taxes on the outflow of local profits accruing to landowners nationwide. The central government taxes and redistributes across places and funds a national public good that depends on inputs produced in various locations. In our framework, exogenous local productivity and residential amenities stand for all dimensions of physical geography.

We give a sufficient statistics characterization of the spatial policies implementing all second-best allocations in this setup. We assume that the government does not observe location preferences. As a result, it is unable to offer type-specific lump-sum transfers and is limited to distortive place-based transfers. Importantly, we also assume that the welfare-enhancing nature (in a Pareto sense) of a potential reform is evaluated in the status quo, in which different workers already live in various places, a criterion we believe bears more political relevance than an ex-ante “veil of ignorance” approach.⁵ Heterogeneity in location preferences opens the door to spatially differentiated welfare effects of spatial policies and to geographic inframarginal winners and losers.⁶ We first provide intuition for the role of location preferences in a two-region example: high enough preference heterogeneity makes the social cost of spatial redistribution encouraging migration too large compared to the agglomeration benefits so that Pareto-improving transfers do not always exist. We also highlight the central role of diminishing returns (e.g., in housing supply) as the key channel through which places gain from being

⁵Precisely, we adopt an ex-post rather than ex-ante Pareto criterion. The ex-ante criterion neutralizes equity concerns by assuming that identical agents draw random preferences and that the planner’s objective is to maximize the common expected utility. The ex-post criterion assumes that reforms are evaluated from observed situations in which heterogeneous agents, already settled, are differently affected by spatial policies. The first approach weights all agents equally, while the second allows the planner to put different weights on different locations in the status quo.

⁶In the standard framework with homogeneous preferences, discrepancies in living standards between cities are arbitrated away by migration pressures so that utility is everywhere equal.

taxed. We show that optimal transfers in the general case should be directed towards places where per capita public spending is higher and away from places able to capture larger per capita amounts of local profits.

We then give a fully computable efficiency test for observed allocations and apply our framework to the French transfer system. Among EU countries, France is the most striking example of large density discrepancies between urban centers and their peripheries and it has been documented that a large share of its GDP is redistributed from high to low density areas (Davezies 2012). We empirically illustrate that the efficiency diagnosis an economist would make strongly depends on the amount of preference heterogeneity she believes led to the observed situation, and that ignoring location preference heterogeneity mistakenly leads to advocating higher net transfers in already dense and rich places. This speaks in favour of seriously accounting for the normative implications of geographic preferences.

Finally, we carry out a revealed preference exercise by empirically investigating the structure of the social welfare weights implied by the current spatial tax and transfer system. Our results suggest that the French planner is roughly utilitarian, but that it further compensates low-density areas in a way that is orthogonal to utilitarianism. This gives support to the idea that the government designs transfers tackling horizontal inequity in access to public goods, which we argue arises from the interplay of unequal density gains from scale economies and heterogeneous location preferences.

We first contribute to the public finance literature on tax competition, public good provision and efficient spatial policies. In a seminal paper, Flatters et al. (1974) provide a formal treatment of efficient population distribution with homogeneous households and locally pure public goods financed by head taxes on residents. Efficient transfers should be set to equalize per capita contributions to local budgets. As for heterogeneous preferences, the authors “argue somewhat heuristically without such an analysis that the basic results derived from the simplest case [...] carry over to this situation” although they carefully highlight the difficulty for the government to implement type-specific taxes and transfers when these preferences are not observed. A rich literature singled out other inefficiencies in local public good provision, such as “rent-sharing” migration externalities when local profits are taxed and partly paid out to residents, and externalities from taxing local profits accruing to residents of other jurisdictions (see Wildasin 1980; Starrett 1980; Boadway and Flatters 1982; Boadway 1982; Watson 1986; Zodrow and Mieszkowski 1986). More recently, Albouy (2012) characterizes efficient

transfers in the presence of public good externalities when the type of heterogeneous households (skill) is observed. The author assumes small if not zero public good scale economies so that spatial transfers predominantly correct the profit taxation migration externality.

More generally, our paper contributes to the urban and economic geography literature on spillovers (Ahlfeldt et al. 2015; Diamond 2016) and on optimal spatial policies (Fajgelbaum and Gaubert 2018; Albouy et al. 2018). We also relate to the literature on spatial misallocation (Desmet and Rossi-Hansberg 2013; Hsieh and Moretti 2015) and misallocation due to specific spatial policies like state or federal income taxes or firm subsidies (Albouy 2009; Fajgelbaum et al. 2015). In particular, Fajgelbaum and Gaubert (2018) characterize efficient transfers in a general framework encompassing production and residential agglomeration externalities that may vary by type (e.g., skill or tastes) and find that the laissez faire is generically inefficient. However, their result crucially hinges on the government’s ability to offer city- and type-specific transfers which evacuates the equity-efficiency trade-off.⁷

We also relate to the fiscal federalism literature examining the optimal balance of powers between central and local governments in the presence of local externalities pioneered by Tiebout (1956) and followed by Oates (1972), Bewley (1981) and Gordon (1983) among others.

This paper contributes to the empirical literature studying the determinants of local public good demand among which population density and urban sprawl. The seminal works of Borcharding and Deacon (1972) and Bergstrom and Goodman (1973) develop simple frameworks to disentangle the various forces driving local public spending. The authors typically find small if not zero net agglomeration gains in public good provision, although their cross-sectional identification, data limitation and incomplete micro-foundation warrants cautious interpretation of the results. A rich cross-sectional literature followed (Ladd 1992, 1994; Carruthers and Ulfarsson 2003; Solé-Ollé and Bosch 2005; Hortas-Rico and Solé-Ollé 2010) which finds negligible gains from more population, but does find that urban sprawl increases public good provision costs. However, Oates (1988) argues that cross-sectional studies severely under-estimate the impact of population because of threshold or “zoo” effects. Indeed, while more populated cities divide public spending among a larger pool of taxpayers and tend to spend less per capita, they also tend to spend more per capita because they reach the critical

⁷In an extension of their framework to unobserved preferences, Fajgelbaum and Gaubert (2018) consider that ex-ante homogeneous workers draw random location preferences and are equally weighted by the government, which evacuates equity concerns by letting the efficient allocation be the one maximizing the common (expected) utility.

mass that makes possible the financing of large indivisible facilities, like zoos. This effect may have researchers conflate public good congestion with the existence of fixed costs, thereby mistaking agglomeration economies for agglomeration dis-economies.

Finally, one can view the normative arguments in this paper as a geographic application of the literature on optimal taxation with discrete occupations (see [Piketty 1997](#); [Saez 2002](#)) and of normative public economics studying the ethics of redistributive policies such as [Fleurbaey and Maniquet \(2011\)](#) or [Saez and Stantcheva \(2016\)](#).

The rest of the paper is organized as follows. Section 2.2 describes the data. Section 2.3 presents new estimates on public good agglomeration economies. Section 2.4 sets up the theoretical framework. Section 2.5 characterizes efficient transfers and provides a theoretical and empirical discussion of the role of location preferences. Section 2.6 presents an empirical exploration into the ethics of spatial redistribution. Section 2.7 concludes.

2.2 Data

We combine a series of rich administrative datasets obtained from various official sources. Detailed description of variable construction will be given in the text when needed. For data on local public spending and government subsidies, we use comprehensive municipal financial accounts (*Balance Comptable des Communes* henceforth BCC, used in [Jannin and Sotura 2019](#)) obtained from the French Ministry of Finance for all 35,000 French municipalities and all 2,000 municipal federations (henceforth MF, which are administrative groupings of close-by municipalities) they belong to, for all years between 2002 and 2016. This rich dataset provides information on various spending items, public assets, local tax revenues, user and business fees as well as various subsidies and transfers received from higher jurisdictions such as *départements* (counties), *régions* (provinces), the government or the EU.

In order to have an in-depth understanding of local tax shares supported by residents, landowners and businesses, we supplement BCC data with administrative local tax data (*Recensement des Éléments d'Imposition à la Fiscalité Directe Locale* henceforth REI) for every year from 2002 to 2016. REI contains data on all tax bases, rates and revenues at the municipal and municipal federation level for local household property and residence taxes, as well as local property and non-property business taxes.

We also use detailed data on natural and urbanized land area at the municipal level for each year between 2008 and 2014 from a government website managed by the French Ministry for

the Environment.⁸ We construct a database on average m^2 housing prices for every even year between 2000 and 2014 at the municipal level, combining housing transactions databases from the notary offices (BIEN database for the Parisian region and PERVAL database for the rest of France) and assume that house transaction prices are the net present value of unobserved rents.

We further exploit an administrative database on municipal population and housing stock (*Fichier des Logements à la Commune* henceforth FILOCOM) available every two years over the period 2000–2014. We use it to construct our yearly municipality-level population measure as well as various useful indicators such as the share of owner-occupiers and alternative measures of urban sprawl. We also obtained access to yearly income tax data at the municipality level over the period 2003–2015 (*Impôt sur le Revenu par Commune* henceforth IRCOM) giving total municipal labor and pension income as well as total national income taxes paid (including tax credits to low-income households and various deductions).

We gather a series of publicly available datasets at the micro and national levels. We use census data to get municipality-level information on socio-demographic characteristics—age profile, education, industry, etc.—for years 1999, 2008 and 2013. We exploit data from the National Statistical Institute (INSEE) on various municipality geography indicators (municipal federation membership, history of municipal boundary changes, etc.). Finally, we use higher-level yearly data on national accounts to measure aggregate capital income (including implicit rental income) to supplement our income tax data.

One objective of this paper is to highlight the purely geographic determinants of unequal access to public goods. We construct a set of relevant variables netted of socio-demographic composition effects using hedonic regression methods. The universe is all mainland municipalities experiencing no boundary change—mergers or separations—between 2002 and 2014, that is, almost all of them barring a few percents. For outcome y_i at the *municipal* level we run

$$\ln y_{it} = X_{it} \cdot \beta_a + \lambda_{jt} + \epsilon_{it} \quad (2.1)$$

where X_{it} is a set of municipal composition characteristics (share of the population at various ages and education levels, share of college-educated executives and share of owner-occupiers), β_a is a metropolitan-area-specific vector of coefficients and λ_{jt} a municipal federation \times year fixed effect. We then use $\hat{\lambda}_{jt}$ and $\hat{\beta}_a$ to predict mean outcome at the municipal federation

⁸See <https://datafoncier.cerema.fr>.

× year level, replacing X_{it} by mean composition at the national level \bar{X}_t . Finally, we take the exponential of this prediction to obtain the composition-neutral outcome. We run this procedure on all relevant intensive variables such as per capita items in financial accounts, per capita local wages, pensions, national and local taxes, m^2 housing consumption — which we then scale back up by population size — and m^2 housing prices.

2.3 Stylized Facts on Public Good Agglomeration Economies

To motivate our theoretical discussion, we present some stylized facts suggesting the existence of scale economies in the consumption of local public goods. Exploiting the panel datasets presented in Section 2.2, we provide new evidence on the relationship between population density and per capita public spending and argue that existing evidence likely understate public good agglomeration economies.

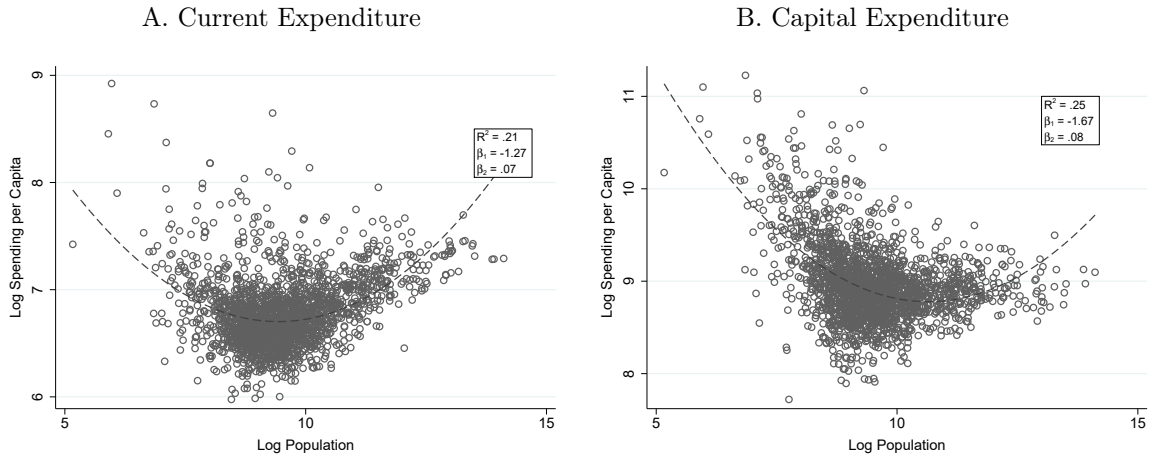
2.3.1 Raw Patterns

In order to focus on scale economies separately from other local externalities and to account for spending by the municipal federation layer on top of municipal spending, we first consolidate all municipality variables at the municipal federation level which leaves us with around 2,000 observations per year.^{9,10} Scale economies in public good consumption imply that public spending should increase less than one-for-one with population, hence we expect to see a decreasing relationship between per capita public spending and population. Figure 2.1 examines this naive cross-sectional relationship in 2014, the last available year. In both Panels, the explanatory variable is log population in the municipal federation. The dependent variable is log current expenditure per capita in Panel A and log capital expenditure per capita in Panel B. Current expenditure are constructed as the sum of annual staff expenditure, maintenance spending, payments for external services and operating subsidies to third parties.

⁹Jannin and Sotura (2019) estimate strong public good spillovers between municipalities within municipal federations.

¹⁰Membership to a municipal federations is not constant during the period covered by our panel. To circumvent this problem, we assign each municipality to its 2015 MF for the whole period when aggregating data at the federation level. This makes the consolidation of municipal and MF financial accounts possible because between 2002 and 2015, MF membership always evolved vertically towards more integration of 2002 blocks of municipalities. That is, isolated municipalities joined existing federations, formed federations with other isolated municipalities and some federations merged between them. No federation split, and almost no municipality changed membership. The underlying assumption is, for example, that an isolated municipality in 2006 already benefited from the total public spending — of both MF and member municipalities — of the MF it joined, say, in 2011.

Figure 2.1: Per Capita Spending vs. Population



Note: These graphs plot log spending per capita against log population across MFs in 2014 and a quadratic fit for which we report the coefficients and the R^2 . Current expenditure are the sum of yearly staff expenditure, maintenance spending, payments for external services and operating subsidies to third parties. Capital expenditure are the current book value of durable facilities and are the sum of all public assets such as schools, transportation infrastructure, parks improvements, sports facilities, museums, art collections, investment subsidies to local clubs, minus the raw value of the land and financial assets.

Capital expenditure—a *stock* position—are defined as the book value of durable facilities and constructed as the sum of all public assets minus the raw value of the land and financial assets. They include schools, transportation infrastructure, parks improvements, sports facilities, museums, art collections, investment subsidies to local clubs, etc.

Both Panels of Figure 2.1 tell a similar story. Spending per capita follows a U-shaped pattern, overall slightly increasing with population for current spending and decreasing with population for capital expenditure. These cross-sectional patterns are reminiscent of results in Ladd (1992) and Ladd (1994). They are in line with overall agglomeration elasticity estimates being close to zero in existing cross-sectional studies, all the more since public capital is never properly accounted for. However, interpreting them through the prism of agglomeration (dis)economies is biased in several ways. First, Oates (1988)’s zoo effect may explain in large parts why spending increases again at higher population levels as the *range* of services widens when cities reach critical population sizes.¹¹ Then, while increasing population on a given urban land area is expected to increase population density and *decrease* the per capita cost of public goods, increasing population holding constant population density is expected to *increase* spending on local public goods as it gets more costly to maintain a given level of public goods

¹¹The needs of populations in larger cities may also systematically differ from those of small city residents.

benefits for new residents at the fringe of the city (see Carruthers and Ulfarsson 2003; Hortas-Rico and Solé-Ollé 2010). Last, local costs and per capita revenues are expected to increase with density which bids up both local prices and demand hence increases per capita public spending.

2.3.2 Descriptive Regressions

Disentangling scale economies from the mechanisms mentioned above would require causally estimating a fully-fledged quantitative model with exogenous migration shocks, while being able to analyse separately the various effects of population changes. Instead, we remain descriptive and provide simple regression evidence. All mechanisms mentioned above are expected to bias the interpretation of Figure 2.1 in the same way by making spending per capita increase again at higher population levels. Below we focus on smaller population variations to absorb discrete changes in the range of provided public goods and explore the role of urban sprawl.

As mentioned, the bundle of local public services is a mixture of yearly services (maintenance, subsidies to associations, staff payroll, etc.) and durable facilities (parks, transports, sports facilities, etc.). We account for this dual structure by assuming a Cobb-Douglas shape for local public goods:

$$\mathbb{G} = \frac{G_s^{1-\psi} G_f^\psi}{N^\kappa} \frac{1}{T^\gamma} \equiv \frac{G}{N^\kappa} \frac{1}{T^\gamma} \quad (2.2)$$

where G_s and G_f are the (unobserved) physical outputs, respectively the quantities of durable facilities and yearly services introduced above, and N is jurisdiction population. Urban sprawl T alters the usefulness of physical outputs at speed γ . Holding constant population size, more residential scatteredness diminishes how much households benefit from given public good outputs. We measure urban sprawl as the km^2 of urbanized land area in the municipality.¹² Parameter $\kappa \in [0, 1]$ is the speed at which public good quantities get congested by population N holding constant the scatteredness of dwellings in the MF, and $1 - \psi$ is the relative importance of durable facilities over annual services.¹³ Assuming a price p_j^G for local public goods, per capita spending is $x_j^G = \frac{p_j^G G_j}{N_j}$.

We first investigate the within-MF relationship between public spending and population.

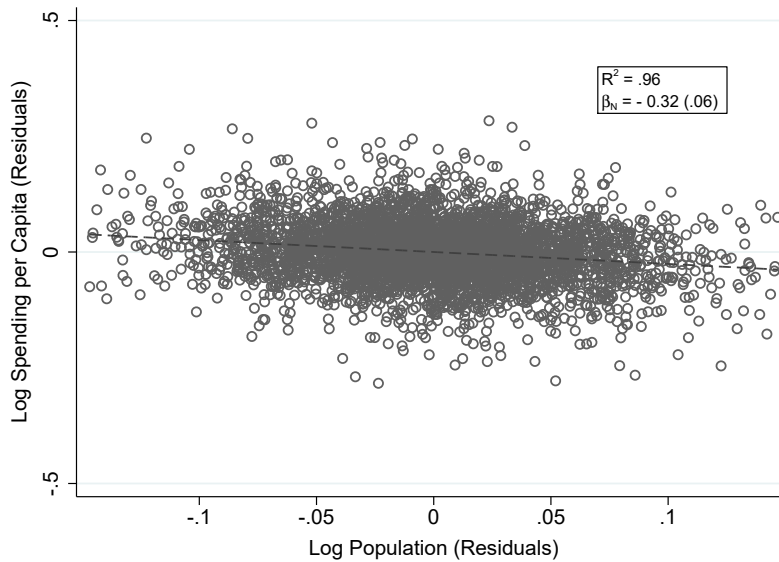
¹²Unlike Hortas-Rico and Solé-Ollé (2010), we control for absolute and not per capita urbanized land area. They are interested in the impact of population density on spending holding constant population size, while we are interested in the impact of population size on spending as measured by parameter κ holding constant the spatial extent of urbanization (see the discussion about density vs. city size in Ahlfeldt and Pietrostefani

This should strongly attenuate the zoo effect as the range of services provided in a MF over our time period is unlikely to evolve dramatically. To mitigate measurement problems raised by looking at time variation (see [Combes and Gobillon 2015](#)) and problems raised by potential short-term rigidity in the supply of local public goods, we look at the longest possible time difference allowed by our data, that is, the 12-year long difference between 2002 and 2014.¹⁴ Unfortunately, data limitations prevents us to also look at changes in urban sprawl over the period 2002–2014. We come back to urban sprawl in a separate specification. We run the following minimal specification aimed at capturing smaller changes in population over time across our 2,000 municipal federations using only extreme years 2002 and 2014:

$$\ln x_{jt}^G = \beta_N \ln N_{jt} + \alpha_j + \lambda_t + \varepsilon_{jt} \quad (2.3)$$

where α_j and λ_t are MF fixed effects and year fixed effects respectively. Figure 2.2 visually inspects the relationship between log per capita spending and log population summarized by

Figure 2.2: Per Capita Spending vs. Density (Within Regressions)



Note: This graph plots the log of our public spending index against log population at the MF level, both variables being first residualized with respect to two-way fixed effects using years 2002 and 2014, and a linear fit. We report the coefficient and R^2 of the two-way fixed effect model (2.3).

2019).

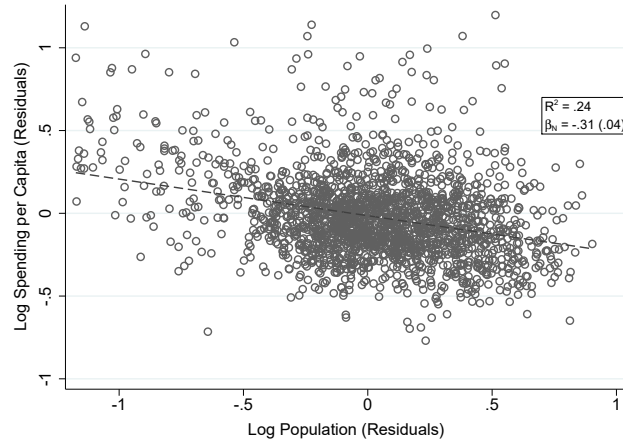
¹³Appendix B.6 strongly supports this constant share assumption with $\hat{\psi} \approx 0.81$.

¹⁴Full rigidity would have per capita spending mechanically decrease one-for-one with population.

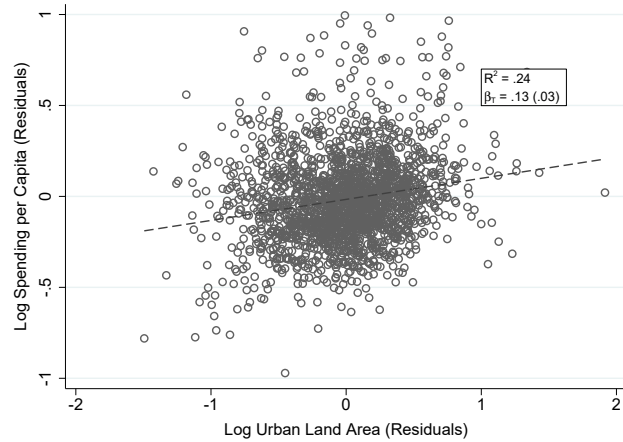
coefficient β_N , that is, after both variables have been residualized with respect to α_j and λ_t . Strikingly, log per capita spending linearly decreases with log population with a reduced-form elasticity of -0.32 (0.06), invalidating the naive interpretation of Figure 2.1.¹⁵

Figure 2.3: Per Capita Spending vs. Density (Cross-Sectional Regressions)

A. Per Capita Spending vs. Population (Constant Urban Land)



B. Per Capita Spending vs. Urban Land Area (Constant Population)



Note: These graphs plot log public expenditure per capita x^G against log population after residualizing both with respect to 100 city size dummies and log urban land area, and log public expenditure per capita against log urban land area after residualizing with respect to 100 city size dummies and log population, with linear fits. Observations are the MFs in 2014. Coefficients and R^2 are that of model (2.4).

¹⁵We also ran the specification using the years 2008 and 2014 and adding log sprawl. As expected, the population coefficient is larger (0.43 ($.07$)) due to the shorter time-span, and that of urban sprawl is zero and insignificant ($.00$ ($.04$)) due to insufficient within-MF variation over time.

As complementary evidence, we run the following specification aimed at capturing density changes in a cross-sectional setting using only the last year 2014 across our 2,000 municipal federations. This time, we look at how per capita spending changes with both population and urbanized land area:

$$\ln x_j^G = \beta_N \ln N_j + \beta_T \ln T_j + \sum_k \lambda_k^N + \varepsilon_j \quad (2.4)$$

where T_j is the urbanized land area in km^2 and the λ_k^N are dummies representing discrete MF population size groups. MF size groups help capture in a crude way discrete changes in the range of provided services, while residual population variation captures proportional scale economies. Results are given in Figure 2.3 for benchmark population steps of 13,500 inhabitants corresponding to 100 groups.¹⁶ The reduced-form elasticity between per capita spending and population is -0.31 (0.04) and close to that of Figure 2.2. Per capita spending is also positively associated with urban sprawl conditional on population, with an elasticity of 0.13 (0.03). Overall, these results show that density is negatively associated with per capita public spending, suggesting scale economies in local public good consumption.

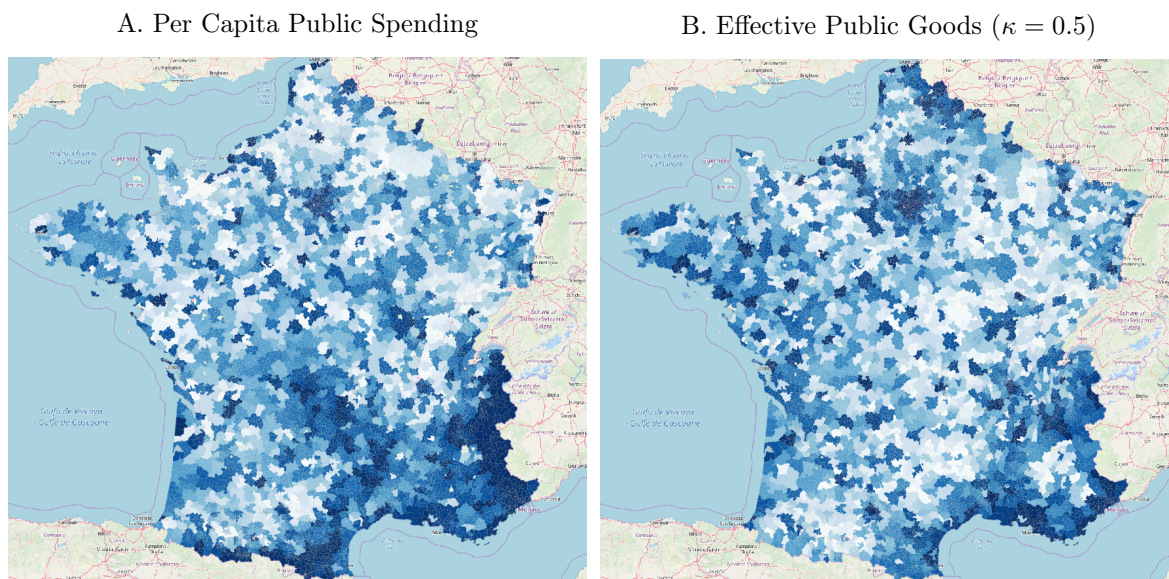
2.3.3 Preliminary Comments on Welfare

Strong public good agglomeration economies have important welfare implications. To guide our intuition on the impact of agglomeration economies on spatial inequities, it is useful to compare the distribution of per capita local public spending and that of local public goods effectively enjoyed by residents. Panel A of Figure 2.4 shows the maps of per capita local public spending x^G , while Panel B captures effective local public good availability obtained by discounting public expenditure by N^κ with an example value for congestion parameter of $\kappa = 0.5$. While per capita public spending is largely concentrated in mountainous and low-density areas, effective public good is much more evenly distributed in this example.

The only remaining pattern seems to be the concentration of effective public goods in large cities that may be explained by large indivisibilities in public goods financing—whose relevance for French municipal federations has been evidenced by Frère and Paty (2011). The apparently even distribution among the other locations suggest that the French spatial redistribution scheme embodies collective preferences for horizontal equity, very present in the

¹⁶Coefficient for urban land area is unchanged as we change the number of groups. As expected, coefficient for population varies between -0.22 (0.04) with 50 groups to -0.42 (0.09) with 500 groups as we further attenuate the zoo effect, and is stable beyond.

Figure 2.4: Per Capita Spending vs. Effective Public Goods



Note: These maps show percentiles of public spending per capita (Panel A) and of effective local public goods (Panel B). Effective public goods are obtained by multiplying per capita spending by $N^{1-\kappa}$ where we take an example value $\kappa = 0.5$.

French political discourse with the concept of “territorial equity” (*Egalité des Territoires*).

Thus, factoring in public good agglomeration economies dramatically changes the perceived spatial distribution of local public good consumption. A fully-fledged economic geography model helps clarify these intuitions in a more systematic way by accounting for efficiency and equity concerns more generally.

2.4 Economic Geography Model with Local Public Goods

Motivated by the evidence of Section 2.3, we develop a Rosen (1974) and Roback (1982) type spatial equilibrium model with endogenous local public goods subject to agglomeration economies. The framework builds on the rich literature studying efficient population distribution (e.g., Flatters et al. 1974; Albouy 2012; Fajgelbaum and Gaubert 2018) and features the — we argue — more realistic assumption that idiosyncratic location preferences are heterogeneous and unobserved. We emphasize that these preferences need not disappear in the long run as they typically capture more than mobility frictions. Since this paper wishes to highlight the spatial nature of efficiency and equity concerns in this context, the framework is kept simple by assuming that workers are homogeneous but for location preferences. Obviously, we could

enrich the model with observed skills. As long as there is unobserved heterogeneity in location preferences conditional on skill, the arguments below carry through.¹⁷

The model works as follows. There are J heterogeneous jurisdictions indexed by j . There is a continuum of workers indexed by i , homogeneous in everything except location preferences. Workers freely choose the city where they both live and inelastically supply one unit of labor. They consume a freely traded good, a non-traded good (e.g., housing) and local non-traded public goods that are subject to agglomeration economies. All these goods are supplied by locally competitive sectors.

Demand for local public goods in each jurisdiction is the outcome of a unanimous vote involving its current residents. Jurisdictions compete between them through residence-based and source-based taxes. All households living in a given jurisdiction are liable to a residence-based head tax. Source-based taxes are property taxes levied on the outflow of local profits accruing to landowners nationwide.

Workers also consume a national public good supplied by the central government. We assume that the latter is a Stackelberg leader, that is, is able to commit to central policies that are determined prior to local governments' and workers' choices. The government taxes and transfers income across locations and supplies the national public good which is produced using inputs from various locations. Sections below describe how the equilibrium is determined in this setup.

2.4.1 Central Government

A national public good $\mathcal{G}(\{\mathcal{O}_j\}_j)$ is supplied by the central government using a combination of local outputs $\mathcal{O}_j = F^{\mathcal{O}}(L_j^{\mathcal{O}})$ produced using local labor. The government makes local hiring decision $L_j^{\mathcal{O}}$ that will be paid the equilibrium local wage w_j , to produce a target level of \mathcal{O}_j in each jurisdiction. Spending on local labor is financed by a general tax and transfer scheme $\{T_j\}_j$ which also corrects migration externalities and redistributes between places in a zero-sum fashion.

Central policies $\{\mathcal{O}_j, T_j\}_j$ are considered fixed by workers and local governments in the spatial equilibrium described below. The government chooses $\{\mathcal{O}_j, T_j\}_j$ by backward induction

¹⁷Figure B.3 in Appendix B.7 further justifies our focus on geography by showing the distribution across municipal federations of a Theil index capturing skill sorting, which is very low in France over the period studied.

among the set of policies that will make the central budget hold in equilibrium:

$$\sum_j N_j T_j = \sum_j L_j^O w_j \quad (2.5)$$

where N_j is population in city j .

2.4.2 Demand for Cities

Worker i chooses to live in the city that maximizes her welfare. Her utility is

$$u_i = \max \{v_{ij}\}_j \quad (2.6)$$

where v_{ij} is the utility level that i gets when living in j . Let's introduce our first structural assumption. We impose that utility break down into a city component v_j indexed by j because of differences between locations in exogenous residential amenities, and an idiosyncratic individual- and city-specific term μ_{ij} that contributes in a multiplicative fashion:

$$v_{ij} = v_j \mu_{ij} \quad (2.7)$$

City utility v_j depends on the consumption of the traded good c_j , the private non-traded good h_j , the local public good \mathbb{G}_j and the national public good \mathcal{G} :

$$v_j = v_j(c_j, h_j, \mathbb{G}_j(G_j, N_j), \mathcal{G}) \quad (2.8)$$

with a slight abuse of notations. Note that (2.8) imposes that all workers in a given city consume the same amount of traded, non-traded and local public goods. Indeed, workers have homogeneous tastes and skills, and the government cannot offer ij -specific transfers since types are unobserved. In addition, local public goods are assumed locally non-excludable so that all residents of j equally enjoy \mathbb{G}_j . We stay general and let $\mathbb{G}_j(G_j, N_j)$ depend on total physical output G_j and population N_j contributing to output congestion. Importantly, we assume that local land policies are such that urban sprawl is fixed over time so that density and population size are equivalent in our framework.

Type- i workers are fully characterized by location preferences $\{\mu_{ij}\}_j$. They are naturally thought of as representing one's attachment for one's city of birth or personalized enjoyment of a city's natural amenities. Section 2.6 discusses how the interpretation of μ_{ij} is central in shaping social preferences. The second structural assumption we make is that the logarithm

of idiosyncratic preferences $\ln(\mu_{ij})$ be i.i.d. across cities and workers and distributed Type-1 Extreme Value $(0, \sigma)$. The share of workers demanding to live in city j upon observing $\{v_k\}_k$ is then

$$N_j = \frac{v_j^{1/\sigma}}{\sum_k v_k^{1/\sigma}} \quad (2.9)$$

where $\sigma \in]0, \infty[$ captures the amount of heterogeneity in location preferences.

2.4.3 Demand for Private Goods

Conditional on living in j , workers

$$\max v_j(c_j, h_j, \mathbf{G}_j(G_j, N_j), \mathcal{G})$$

over $\{c_j, h_j\}$ subject to their budget constraint

$$p c_j + r_j h_j = w_j + \Pi + T_j - \tau_j^h \quad (2.10)$$

where p and r_j are the prices for traded and non-traded goods respectively, w_j is the local wage and Π is total net profits. As detail below, total net profits are redistributed equally to all workers which is why Π has no subscript. T_j is the net per capita transfer from the central government and τ_j^h is the head tax set by the local government, both taken as given by workers. Private optimization yields the usual first-order condition

$$\frac{\partial v_j}{\partial c_j} r_j = \frac{\partial v_j}{\partial h_j} p \quad (2.11)$$

2.4.4 Supply and Ownership

Since there is no individual heterogeneity other than location preferences, all workers living in j receive the same wage w_j .

Supply of the traded good The national good Y_j is locally produced using labor L_j^Y with constant or diminishing returns—because of a fixed factor, e.g., land—and heterogeneous local productivity. Competitive firms' labor demand and traded good supply maximize profits given local wage w_j and national price p :

$$Y_j = F_j^Y(L_j^Y) \quad \text{and} \quad w_j = p \frac{dF_j^Y}{dL_j^Y} \quad (2.12)$$

Housing supply The local non-traded good H_j is produced using labor L_j^H with diminishing returns — because of fixed land — by locally competitive firms maximizing profits given w_j and local housing prices r_j :

$$H_j = F_j^H(L_j^H) \quad \text{and} \quad w_j = r_j \frac{dF_j^H}{dL_j^H} \quad (2.13)$$

Local public good supply Non-traded public good output G_j is locally produced using labor L_j^G with constant or diminishing returns by competitive firms maximizing profits given w_j and local price p_j^G :

$$G_j = F_j^G(L_j^G) \quad \text{and} \quad w_j = p_j^G \frac{dF_j^G}{dL_j^G} \quad (2.14)$$

Ownership structure We assume that land is the residual claimant of local profits and that land ownership is evenly distributed among households. All households hence own the same fraction of total net profits. Total net profits are the sum of local profits accruing to land ownership net of local property taxes:

$$\Pi = \sum_j \Pi_j (1 - \tau_j^p) \quad (2.15)$$

where

$$\Pi_j = p Y_j + r_j H_j + p_j^G G_j - w_j (L_j^Y + L_j^H + L_j^G) \quad (2.16)$$

is the outflow of local gross profits accruing to landowners nationally and τ_j^p is the local property tax.

Market clearing We complement equilibrium equations with market clearing conditions. Local labor market must clear:

$$L_j^Y + L_j^H + L_j^G + L_j^O = N_j \quad (2.17)$$

where the elements of $L_j \equiv \{L_j^Y, L_j^H, L_j^G, L_j^O\}$ are defined above. The condition for local housing market clearing is

$$H_j = N_j h_j \quad (2.18)$$

Finally, the market for the traded good clears:

$$\sum_j Y_j = \sum_j N_j c_j \quad (2.19)$$

2.4.5 Demand for Public Goods and Tax Competition

Demand for local public goods in j is set by the local government through a unanimous vote on a head tax τ_j^h imposed on all residents and a property tax τ_j^p imposed on the outflow of local profits accruing to landowners nationwide.¹⁸ We assume that local governments take part in a Cournot-Nash competition, i.e., set the head tax and the property tax taking other jurisdictions' policies as given.¹⁹ Local governments also take the central government's policies $\{\mathcal{O}_j, T_j\}_j$ as given.

Importantly, we assume that local governments are non-myopic, i.e., perfectly anticipate migration responses to own policies. Because migration responses impact local labor and goods markets, jurisdictions also foresee these changes. Being small, jurisdictions however do not anticipate how the price of the traded good p adjusts in equilibrium. A local jurisdiction maximizes the welfare of its current residents $v_j(c_j, h_j, \mathbb{G}_j(G_j, N_j), \mathcal{G})$ over $\{G_j, \tau_j^h, \tau_j^p\}$ subject to equations (2.8) to (2.18) and local budget constraint

$$p_j^G G_j = \tau_j^h N_j + \tau_j^p \Pi_j \quad (2.20)$$

taking as given p , other jurisdictions' policies $\{\tau_k^h, \tau_k^p\}_{k \neq j}$ and central policies $\{\mathcal{O}_k, T_k\}_k$. Although migration anticipation complicates the analysis, it ensures that local governments provide public goods efficiently from a social perspective by respecting a local Samuelson rule.

Proposition 1. *Each jurisdiction satisfies a Samuelson rule as an interior condition for τ_j^h :*

$$\frac{p_j^G}{p} \frac{\partial v_j}{\partial c_j} = N_j \frac{\partial \mathbb{G}_j}{\partial G_j} \frac{\partial v_j}{\partial \mathbb{G}_j} \quad (2.21)$$

Proof. See Appendix B.1. □

This Samuelson rule accounts for congestion through the term $\frac{\partial \mathbb{G}_j}{\partial G_j}$ which captures how a marginal change in public output impacts local public goods. For example, when $\mathbb{G}_j(G_j, N_j) = \frac{G_j}{N_j^\kappa}$ with κ the congestion parameter, an extra unit of public good output is discounted by N_j^κ before being equally enjoyed by all residents, and the Samuelson rule becomes $\frac{p_j^G}{p} \frac{\partial v_j}{\partial c_j} =$

¹⁸Head taxes can be implemented through a variety of residence-based tax instruments (e.g., on wages and profits earned from land and capital property nation-wide) since households are homogeneous but for location preferences. This is true as long as these taxes are non-distorting conditional on residence (e.g., when labor supply is inelastic). We hence rule out taxes levied on consumption such as taxes proportional to housing as they would create local distortions that would require central intervention.

¹⁹We assume that the strategic instruments are taxes and that spending adjusts endogenously.

$N_j^{1-\kappa} \frac{\partial v_j}{\partial G_j}$. When $\kappa = 0$, it collapses to the familiar Samuelson rule $\frac{p_j^G}{p} \frac{\partial v_j}{\partial c_j} = N_j \frac{\partial v_j}{\partial G_j}$ stating that when public goods are fully shared, the marginal cost for a worker of providing an extra unit of public good should equate the sum of marginal benefits to all workers in the jurisdiction.

For completeness, we also report the first-order condition for the property tax τ_j^p :

$$\Pi_j \left(\frac{1}{p_j^G} \frac{\partial v_j}{\partial G_j} \frac{\partial G_j}{\partial G_j} - \frac{\partial v_j}{\partial c_j} \right) \left(1 + \frac{\partial \ln v_j}{\partial \ln N_j} \Big|_{\tau} \times B_j \right) + C_j \sum_{k \neq j} \Pi_k N_k \frac{v_j}{v_k} \frac{\partial v_k}{\partial c_k} = 0 \quad (2.22)$$

where B_j and C_j are defined in Appendix B.1. Appendix B.1 shows that a myopic jurisdiction would violate the Samuelson rule when allowed to levy both head taxes and profit taxes. Taxation would then be socially inefficient because myopic jurisdictions would fund part of their public spending through profits accruing to landowners nationwide without internalizing the cost to non-residents (Starrett 1980). This externality disappears when local governments correctly anticipate migration responses since the prospect of additional congestion from new residents deters excessive taxation of non-residents (Boadway 1982). Assuming that jurisdictions anticipate migration responses hence allows us to evacuate the profit taxation externality and to limit the scope of the efficiency discussion to population distribution and migration externalities only.

2.4.6 Equilibrium

Given central government policies $\{\mathcal{O}_j, T_j\}_j$, a spatial equilibrium is defined by quantities $\{N_j, c_j, h_j, Y_j, H_j, G_j, L_j, \Pi_j\}_j$, utility levels $\{v_j\}_j$, local policies $\{\tau_j^h, \tau_j^p\}_j$, local prices $\{r_j, w_j, p_j^G\}_j$ and national price p such that population distribution satisfies the free mobility condition (2.9) with local utilities defined by (2.8); demand for the traded good and housing is set by (2.10) and (2.11) with profits defined by (2.16); supply of the traded good, housing and local public goods as well as labor demand in these three sectors are given by (2.12), (2.13) and (2.14); local public good demand and local taxes are given by (2.20), (2.21) and (2.22); and all markets clear according to (2.17), (2.18) and (2.19). The set of feasible $\{\mathcal{O}_j, T_j\}_j$ is determined by the central government budget constraint (2.5).

Given the existence of agglomeration forces in our model, the equilibrium may in not be unique. A sufficient condition for existence and uniqueness of a laissez-faire ($T_j = 0$) equilibrium in parameterized applications of our framework is that dispersion forces—location preference heterogeneity and diminishing returns to scale in production technologies—be

stronger than public good scales economies (Redding 2016). As shown in Fajgelbaum and Gaubert (2018), such conditions also ensure the existence and uniqueness of the equilibrium at the optimal spatial policy.

2.5 Optimal Policies

2.5.1 Intuition in a Two-Region Example

We illustrate the mechanism through which heterogeneity in location preferences, by slowing down migration responses, increases the social cost of migration incentives (spatial taxes and transfers) relatively to their social benefits (agglomeration gains) such that improving welfare in a Pareto sense over the laissez-faire economy may be impossible. We also highlight the role of diminishing returns, for instance in housing supply, as a necessary channel for Pareto improvements in the presence of agglomeration externalities and Pigovian transfers. To this end, we work with a simple application of our framework with two regions and Cobb-Douglas preferences. Detailed steps are given in Appendix B.2.

We assume that technology for traded and local public goods exhibits constant returns to scale. Geography only enters the model through exogenous local productivity $z_j^Y = z_j^G = z_j$ which is constant in both sectors but varies across space with $z_1 > z_2$. Labor demand yields $w_1 = z_1 > w_2 = z_2$. Technology for housing is uniform across space and exhibits diminishing returns with constant housing supply elasticity η . Agents have utility $v_{ij} = (c_j^{1-\alpha} h_j^\alpha)^{1-\phi} (\mathbb{G}_j)^\phi \mu_{ij}$ where $\{\alpha, \phi\}$ are taste parameters in $]0, 1[$. For simplicity, we abstract from the national public good and assume that local public goods are financed by the residence tax only, i.e., $\mathbb{G}_j = \tau_j^h N_j^{1-\kappa}$. Public good benefits hence increase with population with constant elasticity $1 - \kappa > 0$. Combining local demand and supply equations, one shows that

$$v_j = (z_j + T_j + \Pi)^{1-\epsilon_H} (N_j)^{\epsilon_G - \epsilon_H} \quad \text{with} \quad \epsilon_H = \frac{\alpha(1-\phi)}{1+\eta}, \quad \epsilon_G = (1-\kappa)\phi \quad (2.23)$$

where $0 < \epsilon_H < 1$ and $0 \leq \epsilon_G < 1$ are the land congestion elasticity and the public good agglomeration elasticity respectively. Utility increases with disposable income, but less than one for one as more income means higher housing prices. Holding profits constant, utility increases or decreases with region population depending on the relative strength of public good agglomeration and land congestion forces. We assume that dispersion forces overall are

stronger than agglomeration forces:

$$\sigma + \epsilon_H > \epsilon_G \quad (2.24)$$

to ensure that the equilibrium is unique in the laissez-faire economy and that the planner's objective is concave. Note that heterogeneity in location preferences $\sigma > 0$ implies that one may have $\epsilon_G > \epsilon_H$, i.e., public good agglomeration forces may be stronger than land congestion forces.

Small reform approach Central policies $\{T_1, T_2\}$ must balance the budget, i.e., $N_1 T_1 + N_2 T_2 = 0$. Let us introduce a small transfer $dT_1 > 0$ starting from the laissez faire ($T_1 = T_2 = 0$) while T_2 adjusts with $dT_2 < 0$. Only households that are initially marginally indifferent between 1 and 2 migrate. Using an envelope argument, welfare of these marginal households is not affected to a first order. Utility of infra-marginal residents in each region changes according to:

$$\begin{aligned} \frac{d \ln v_1}{dT_1} &\propto \frac{1}{z_1 + \Pi} + \frac{1}{z_1 + \Pi} \frac{d\Pi}{dT_1} + \frac{\epsilon_G - \epsilon_H}{1 - \epsilon_H} \frac{1}{N_1} \frac{dN_1}{dT_1} \\ \frac{d \ln v_2}{dT_1} &\propto -\frac{1}{z_2 + \Pi} \frac{N_1}{N_2} + \frac{1}{z_2 + \Pi} \frac{d\Pi}{dT_1} - \frac{\epsilon_G - \epsilon_H}{1 - \epsilon_H} \frac{1}{N_2} \frac{dN_1}{dT_1} \end{aligned} \quad (2.25)$$

In each equation, the first term is the direct impact of dT_1 on infra-marginal households' budget, positive in region 1 and negative in region 2. The second term is the change in profits, positive because households are enticed to locate in the more populated region 1 where housing prices increase faster than they decrease in 2.²⁰ The third term is the gain from agglomeration net of extra land congestion, negative or positive depending on $\epsilon_G - \epsilon_H$ but with opposite signs in region 1 and region 2. To know the sign of welfare changes in each region, we plug the expressions for $d\Pi$ and dN_1 in (2.25).

Proposition 2. *The sign of welfare change in each region at $T_1 = T_2 = 0$ is*

$$\begin{aligned} \text{sign} \left\{ \frac{dv_1}{dT_1} \right\} &= \text{sign} \left\{ N_1 \frac{z_1 - z_2}{z_2 + \Pi} \epsilon_G + \sigma \right\} \\ \text{sign} \left\{ \frac{dv_2}{dT_1} \right\} &= \text{sign} \left\{ N_2 \frac{z_1 - z_2}{z_1 + \Pi} \epsilon_G - \sigma \right\} \end{aligned} \quad (2.26)$$

Benchmark with no externality or no geography Absent public good agglomeration gains, i.e., $\epsilon_G = 0$, welfare changes have opposite signs (that of σ and $-\sigma$), consequently no Pareto improvement is feasible. As expected, the laissez-faire economy is efficient in the

²⁰From an optimal policy viewpoint, accounting for housing profits is necessary as land congestion is not an externality.

absence of externality. When dispersion in location preferences vanishes, that is, $\sigma \rightarrow 0$, the spatial equilibrium equalizes utilities between 1 and 2. The small tax has zero impact on welfare since taxes and transfers are exactly compensated for by changes in local housing prices and profits through migration responses. When $\sigma > 0$, dT_1 has a negative effect on region 2 since out-migration is too small to compensate it through decreased housing prices and increased profits, and conversely has a positive effect on region 1.

Absent geography ($z_1 = z_2$) welfare changes have the sign of σ and $-\sigma$ so that no Pareto improvement is feasible either. In this case $N_1 = N_2$ so $dT_1 = -dT_2$, agglomeration gains in 1 equal agglomeration losses in 2, changes in housing prices in 1 equal changes in housing prices in 2 and $d\Pi = 0$ so that there are no aggregate gains. When $\sigma \rightarrow 0$, utility must be equalized between 1 and 2 hence stays constant in the absence of aggregate gains. When $\sigma > 0$, the net effect is positive for region 1 and negative for region 2 since migration responses are smaller.²¹

Thus, when agglomeration externalities have constant elasticity, it is the interaction between agglomeration gains and differences in geographic fundamentals that calls for policies correcting population distribution.

Cases with geography and agglomeration \leq congestion In this case $0 < \epsilon_G \leq \epsilon_H$. One can have $\sigma \rightarrow 0$ while satisfying (2.24). When $\sigma \rightarrow 0$, welfare changes are positive at the limit in both cities since $z_1 > z_2$ and there are now aggregate gains to giving incentives to migrate to 1. It is Pareto-improving to introduce a small subsidy $dT_1 > 0$ and the small tax $dT_2 < 0$ that balances the budget. This is the usual setup examined in the literature. By continuity, Pareto improvements are feasible at least for small $\sigma > 0$. Since welfare change in region 1 is always strictly positive, the only Pareto improving candidate policy is $dT_1 > 0$. One can find a sufficient condition for welfare change in region 2 to always be strictly negative, that is, a condition under which no Pareto improving reform exists, featuring only structural parameters and the model's geography.

Proposition 3. *When $0 < \epsilon_G \leq \epsilon_H$, there is a non-empty region for $\sigma > 0$ in which Pareto improving transfers exist. However, when heterogeneity in location preferences is high enough, no policy can achieve Pareto gains. A sufficient condition for this is*

$$\sigma > \epsilon_G \frac{z_1 - z_2}{z_1} \tag{2.27}$$

²¹Here we assumed $\epsilon_G \leq \epsilon_H$ such that net gains from density are negative, to allow for the possibility that $\sigma \rightarrow 0$.

When location preferences are sufficiently heterogeneous, migration pressures are too weak to ensure that housing prices and profits adjust to have direct payers indirectly benefit from the transfer scheme.

Cases with geography and agglomeration > congestion In this case $\epsilon_G > \epsilon_H$ and one can allow a minimum limit of $\sigma \rightarrow \epsilon_G - \epsilon_H$ to satisfy (2.24). From Proposition 3 we know that there is a small region where $\epsilon_G > \epsilon_H$ for which Pareto improvements are feasible when $\sigma \rightarrow \epsilon_G - \epsilon_H$. This is because although net migration gains in region 2 are strictly negative, region 2 is still compensated through the increase in profits Π . When σ increases too much no Pareto improving reform is feasible as stated by Proposition 3. Finally, one can show that no welfare gains are feasible when agglomeration forces are too large compared to land congestion forces.

Proposition 4. *A sufficient condition for the impossibility of any Pareto improving reform whatever the value of $\sigma > \epsilon_G - \epsilon_H$ is that the agglomeration force is too strong relative to the land congestion force:*

$$\epsilon_G > \frac{z_1}{z_2} \epsilon_H \quad (2.28)$$

Proposition 4 highlights the central role of the housing market as the channel by which the taxed region 2 gets compensated by lower housing prices and higher profits in this example. For example, with perfectly elastic housing supply ($\eta \rightarrow \infty$ so $\epsilon_H \rightarrow 0$) no Pareto improving reform exists in this setup.

Armed with these intuitions, we now characterize the efficiency frontier in the general case by the way of sufficient observable statistics that are robust to the model's primitives.

2.5.2 Efficient Allocations

We come back to the general model of Section 2.4. A standard method to find all Pareto efficient allocations is to maximize the social welfare function

$$W = \mathbb{E} [\omega_i \ln u_i] \quad (2.29)$$

for all combinations of individual Pareto weights $\{\omega_i\}_i$ where $\omega_i > 0$, subject to resource and market clearing constraints.²² Each vector of Pareto weights $\{\omega_i\}_i$ (and its multiples) is

²²That is, if the planner's problem is convex. We make the assumption that underlying dispersion forces—e.g., the housing market and location preferences—are strong enough to ensure convexity of the planner's problem.

associated with a Pareto efficient allocation and pins down a particular equity objective on the efficiency frontier.

Reaching first-best efficiency would require the planner to be able to freely displace populations and offer ij -specific consumption bundles. We impose additional constraints which will make efficiency only second-best. First, we assume that the planner does not observe types i and has to give the same c_j and h_j to all individuals living in j . Together with the assumption that G_j is locally non-excludable, it implies that the same v_j has to be granted to all residents of j . Thus, the planner is bound to optimize over $\{v_j, N_j, c_j, h_j, G_j, H_j, L_j, \mathcal{O}_j\}_j$. Second, the planner is constrained by agents' free location decisions:

$$u_i = \max \{v_j(c_j, h_j, \mathbb{G}_j(G_j, N_j), \mathcal{G}) \mu_{ij}\}_j$$

Choice of Pareto weights As in most optimal taxation settings, the planner may choose Pareto weights *in abstracto* based only on the knowledge of the distribution of the μ_{ij} 's even though individual realisations are unobserved. This would require the planner to know where each type will locate in equilibrium even though type is unobserved (and should stay unobserved after the equilibrium is realised since several types will end up in the same location). Alternatively, the planner may choose Pareto weights as a function of individuals' observed situation in the status quo in the spirit of [Saez and Stantcheva \(2016\)](#).²³ In this case, Pareto weights would realistically be constrained to be homogeneous across households that live in the same city in the status quo. The assumption about how weights are chosen does not affect the efficiency discussion below.

Proposition 5. (Envelope Theorem) *Given Pareto weights $\{\omega_i\}_i$, any small deviation from the status quo has the following effect on social welfare:*

$$dW = \sum_j N_j \Omega_j \frac{dv_j}{v_j} \quad (2.30)$$

where

$$\Omega_j \equiv \frac{1}{N_j} \mathbb{E} \left[\omega_i \mathbf{1}_{[v_{ij} = \max\{v_{ik}\}_k]} \right] \quad (2.31)$$

is the average Pareto weight across households living in city j in the status quo.

Proof. We give intuition for this discrete choice version of the envelope theorem. Infra-marginal residents of j , weighted Ω_j on average, benefit from the first-order change in local conditions

²³However, weights are kept exogenous in the maximization procedure, a necessary condition for non-violation of the Pareto Principle.

v_j . Only marginal residents migrate in response to a small change in the v_j 's. Because they are initially indifferent between two or more locations, migration does not have a first order effect on their own welfare holding constant the v_j 's. However, migrants will also benefit from small changes in the v_j 's. Because both these changes and the size of migrants are marginal, the compound effect on W is only of second order. \square

The status quo $\{v_j, N_j, c_j, h_j, G_j, H_j, L_j, O_j\}_j$ is efficient if and only if it is a critical point for social welfare W for some profile of positive weights $\{\omega_i\}_i$. We report below two salient conditions.

Proposition 6. (Efficient Allocation) *The optimal population allocation satisfies*

$$\frac{dF_j^Y}{dL_j^Y} + N_j \frac{\partial v_j / \partial G_j}{\partial v_j / \partial c_j} \frac{\partial G_j}{\partial N_j} = c_j + h_j \frac{\partial v_j / \partial h_j}{\partial v_j / \partial c_j} + E + (E_j - E) \quad \text{for all } j \quad (2.32)$$

where $E = \sum_j N_j E_j$. Local output should be produced in each location such that

$$\sum_k N_k \frac{\partial v_k / \partial \mathcal{G}}{\partial v_k / \partial c_k} \frac{\partial \mathcal{G}}{\partial O_j} = \frac{\partial F_j^Y / \partial L_j^Y}{\partial F_j^O / \partial L_j^O} \quad \text{for all } j \quad (2.33)$$

Proof. See Appendix B.3. \square

Equation (2.33) simply states that national public goods should be produced according to Samuelson rules with no consideration for equity nor migration externalities. This notably evacuates indirect redistribution through central public demand stimulating local labor markets. Equation (2.32) characterizes the efficient population distribution. The first term of the left-hand side is the benefit of moving a worker to location j through increased production. The second term is the congestion of local public goods brought by an extra worker expressed in money terms. The first term of the right-hand side is the cost of moving one worker to location j as this worker now needs to consume c_j and h_j . The second term E captures the cost of an additional worker in the economy.

The novelty in (2.32) is the third term $E_j - E$ which is the social opportunity cost of having one more worker located in j and not elsewhere. This cost is increasing in the average welfare weight of all workers relative to workers of city j . It reflects the fact that moving one worker to j implies taxing residents of all other cities in order to create the migration incentive that must hold in a decentralized equilibrium.

2.5.3 Optimal Transfers

We now characterize the central government's tax and transfer system that decentralizes the planner's efficient allocation, expressed in terms of estimable sufficient statistics (see [Chetty 2009](#)). While the public finance literature has traditionally expressed optimal income taxes with sufficient statistics, optimal Pigouvian taxes may also be characterized in this way—although the estimation requirements are higher (see [Kleven 2018b](#) for a recent discussion on sufficient statistics).²⁴ Importantly, all endogenous variables below are given at the optimal policy so this characterization is implicit.

Proposition 7. (Optimal Transfers) *The optimal per capita transfer T_j is*

$$T_j = T + T_j^G + T_j^\Pi + T_j^\Omega \quad (2.34)$$

where

$$\begin{aligned} T &= \sum_j N_j T_j = \sum_j L_j^\mathcal{O} w_j \\ T_j^G &= x_j^G (1 + \epsilon_{G_j, N_j}) - \sum_k N_k x_k^G (1 + \epsilon_{G_k, N_k}) \\ T_j^\Pi &= - \left[x_j^G (1 - \tau_j) - \sum_k N_k x_k^G (1 - \tau_k) \right] \\ T^\Omega &= \sigma (\Omega_j - \lambda_j v_j) \end{aligned} \quad (2.35)$$

Proof. Plug equilibrium equations into (2.32) and the first-order condition for v_j . \square

Proposition 7 characterizes the transfers achieving the government's efficiency and equity objectives. The first term T in (2.34) funds the national public good in a lump-sum, i.e., geography-neutral fashion.

The second term T_j^G is the corrective tax and transfer scheme that tackles public good agglomeration externalities, with $x_j^G = \frac{p_j^G G_j}{N_j}$ the per capita spending on local public goods and $\epsilon_{G_j, N_j} = \frac{\partial G_j}{\partial N_j} \frac{N_j}{G_j}$ the congestion elasticity of local public goods. With the efficient public good provision result of Proposition 1, agglomeration elasticities $\{1 + \epsilon_{G_j, N_j}\}_j$, per capita public spending $\{x_j^G\}_j$ and population distribution $\{N_j\}_j$ are the sufficient statistics characterizing

²⁴“Sufficient statistics” optimal policy formulas typically feature reduced-form elasticities capturing how the behavior causing the externality—here residential choice—changes with the policy trying to correct it. Our formulas do not because we put some structure on the shape of migration responses, which allows to drastically reduce the parameter space. Had we not, formulas would feature various reduced-form city population elasticities to tax and transfer changes, including around 2,000 cross-elasticities. Our approach is hence halfway through classic public finance sufficient statistics studies (since our optimal transfers dependent on a reduced set of observables and parameters) and more structural approaches.

the efficient tax.²⁵ Assuming constant agglomeration elasticity $1 + \epsilon_{G,N} = 1 - \kappa$ for example, the corrective tax becomes $T_j^G = (1 - \kappa)(x_j^G - \bar{x}^G)$ and is zero only if all locations spend the same amount per capita on local public goods. The pigovian net transfer should be higher in municipalities where per capita public spending is higher. Other things being equal, this typically means transferring resources towards high-productivity and low-density places.

The third term T_j^Π tackles rent-sharing externalities, with $\tau_j = \frac{\tau_j^h N_j}{p_j^G G_j}$ the share of local spending paid for by residents and $1 - \tau_j = \frac{\tau_j^p \Pi_j}{p_j^G G_j}$ the share paid for by local profit taxation. When an individual migrates to location j , she automatically gets a fraction of $\tau_j^p \Pi_j$ —the share of local profits captured by residents of j through the property tax—which was previously allocated to current residents (see [Boadway and Flatters 1982](#)). This creates a negative migration externality that vanishes when property taxation disappears. The sufficient statistics characterizing the efficient tax are per capita public spending $\{x_j^G\}_j$, residents' tax share $\{\tau_j\}_j$ and population distribution $\{N_j\}_j$. Transfers correcting rent-sharing externalities should tax away all profits captured by local governments.

The fourth term T_j^Ω achieves equity goals by redistributing between locations, with Ω_j the average Pareto weight in j and λ_j a multiplier. Under homogeneous preferences ($\sigma \rightarrow 0$), because the term in brackets is typically bounded, the redistributive tax goes to zero. Indeed, the spatial equilibrium guarantees that utilities are everywhere equal when $\sigma \rightarrow 0$ so that equity concerns vanish. As mentioned, the introduction of observed skill groups would only generalize the model leaving the above argument intact. Unobserved location preferences—when not fully correlated with skill—open the door to geographic equity concerns orthogonal to what skill-based transfers can compensate.

Finally, note that corrective and redistributive tax and transfer schemes are self-financed at the optimal policy:

$$\sum_j N_j T_j^G = \sum_j N_j T_j^\Pi = \sum_j N_j T_j^\Omega = 0$$

Proposition 7 implies that the second welfare theorem breaks down. Unsurprisingly, the first two Pigovian taxes T_j^G and T_j^Π dealing with migration externalities are place-based policies. Importantly and unlike in [Flatters et al. \(1974\)](#), [Albouy \(2012\)](#) or [Fajgelbaum and Gaubert \(2018\)](#), the third tax T_j^Ω related to equity goals is not people-based but place-based,

²⁵Without loss of generality, we assumed $\frac{\partial G_j}{\partial G_j} \frac{G_j}{G_j} = 1$ to ease exposition.

because individual types — here idiosyncratic location preferences — are unobserved. Because redistribution can only be achieved by targeting locations, compensating losers of a reform targeting population efficiency can only be done in a spatially distorting way.

2.5.4 An Efficiency Test

The above sections characterized the efficient transfers given a set of Pareto weights $\{\omega_i\}_i$. In this section, we invert that logic and derive a simple efficiency test for observed allocations. Given any allocation, we reverse-engineer the weights that rationalize it as one maximizing social welfare (2.29). Assuming homogeneity of degree 1 for v_j , one easily shows that $\lambda_j v_j = x_j$ where x_j is total per capita spending (public and private) in j .

Proposition 8. (Implied City Weights) *Assume homogeneity of degree 1 for v_j . The city weights that make government policies $\{T_j, \mathcal{O}_j\}_j$ a critical point for social welfare W are*

$$\begin{aligned} \tilde{\Omega}_j &= x_j \\ &+ \frac{1}{\sigma} \left(T_j - \sum_k N_k T_k \right) \\ &- \frac{1}{\sigma} \left(x_j^G (1 + \epsilon_{G_j, N_j}) - \sum_k N_k x_k^G (1 + \epsilon_{G_k, N_k}) \right) \\ &+ \frac{1}{\sigma} \left(x_j^G (1 - \tau_j) - \sum_k N_k x_k^G (1 - \tau_k) \right) \end{aligned} \tag{2.36}$$

This follows directly by reversing Proposition 7. The assumption of homogeneity of degree 1 makes the implied weights fully computable given a value for σ . Proposition 8 characterizes the set of weights that the government must assign to households in the status quo to justify the absence of further policy intervention according to objective (2.29). It is stated in terms of average city weights in the status quo. In coherence with the fact the government may in practice only condition weights on observed city of residence, a feasible set of individual weights is when all agents living in a same city j in the status quo equilibrium are given the same weight $\tilde{\Omega}_j$.

One can use Proposition 8 to diagnose the inefficiency of various situations and relate to the existing literature. Indeed, when the planner's problem is convex, transfers are efficient if and only if they maximize (2.29) for a given set of Pareto — that is, strictly positive — weights. This means that observed central policies are efficient only if they imply city weights that are Pareto weights, i.e., only if $\tilde{\Omega}_j > 0$ for all j . We use proposition 8 to revisit two recent results

in the optimal spatial policies literature.

Proposition 9. (Albouy (2009) with Location Preferences) *When workers have heterogeneous location preferences, efficient spatial transfers need not compensate the “unequal geographic burden” of national taxes funding national public goods.*

In a seminal paper, [Albouy \(2009\)](#) shows that national income taxes, although place-independent, are made de facto place-dependent because wages are partly determined by location. Under homogeneous preferences, this geographic tax gap creates a deadweight loss. To see how this result changes with heterogeneous preferences, call the spatial laissez faire the situation in which the central government only supplies the national public good according to (2.33), financed through a geography-neutral head tax T . For simplicity, let’s assume away agglomeration and rent-sharing externalities. In this case, implied average city weights in the laissez-faire economy are $\tilde{\Omega}_j = x_j > 0$ where x_j is total per capita spending in j in the equilibrium arising under the national lump-sum policy T . Unsurprisingly, the spatial laissez faire can be related to a set of Pareto weights—since implied average city weights are strictly positive—so it is efficient. It may however not be considered equitable. The government may deviate from the head tax and carry out spatial redistribution while keeping the economy on the efficiency frontier as long as both (2.33) and $x_j + \frac{1}{\sigma} (T_j - T) > 0$ hold for all j . By continuity and since $x_j > 0$ in the laissez-faire economy, introducing some geographic redistribution from $T_j = T$ will not break efficiency. Hence, unlike in [Albouy \(2009\)](#) or [Albouy \(2012\)](#), efficient policies here need not compensate the “unequal geographic burden” of national income taxes funding national public goods. Unobserved location preferences introduce spatial equity concerns and the (second-best) efficient funding of national public goods may feature spatially differentiated contributions.

Proposition 10. (Fajgelbaum and Gaubert (2018) with Location Preferences) *Under homogeneous location preferences, the laissez-faire economy is generally inefficient. When heterogeneity in location preferences is high enough, the laissez-faire economy is efficient.*

In the presence of agglomeration gains and rent-sharing externalities, inefficiency of the spatial laissez faire depends on location preference heterogeneity. Under homogeneous preferences ($\sigma \rightarrow 0$), efficiency requires $x_j^G (\epsilon_{G_j, N_j} + \tau_j) - \sum_k N_k x_k^G (\epsilon_{G_k, N_k} + \tau_k) = 0$ for all j , which is impossible unless the term $x_j^G (\epsilon_{G_j, N_j} + \tau_j)$ is identical across space. Barring knife-edge cases (e.g., constant agglomeration elasticity and no differences in geographic fundamentals), the spatial laissez faire is never efficient under homogeneous location preferences. This nests the results in [Flatters et al. \(1974\)](#), [Albouy \(2012\)](#) or [Fajgelbaum and Gaubert \(2018\)](#) for instance.

However, with large enough preference heterogeneity, Pareto-improving spatial policies become unfeasible. When $\sigma \rightarrow \infty$, the last two terms of Proposition 8 will typically go to zero.²⁶ The spatial laissez faire ($T_j = T$) will correspond to a set of Pareto weights and be efficient. Put differently, the government cannot improve everyone's welfare through place-based transfers when preference heterogeneity is too high, even in the presence of migration externalities, because spatial policies are bound to hurt infra-marginal residents somewhere. Higher preference heterogeneity dampens migration responses, which may make the social redistribution cost of migration incentives larger than the social benefits of agglomeration gains.

2.5.5 Efficiency of Observed Transfers

We take Proposition 8 to the data and use it to diagnose the inefficiency of observed spatial transfers in France. As mentioned, the existence of some negative implied weights is a sufficient condition for the observed allocation to be inefficient. Under our assumption that preference draws are i.i.d. across cities and workers and distributed Type-1 Extreme Value, these weights also give the direction of Pareto-improving marginal reforms.

Proposition 11. (Pareto-Improving Marginal Transfers) *Given a status quo equilibrium with transfers, Pareto-improving marginal transfers should be positive in locations having negative implied weights, and negative in locations having positive implied weights.*

Proof. See Appendix. □

Construction of variables Using our rich administrative data, we are able to construct the variables of Proposition 8. Per capita spending on private and public goods $x_j = w_j + \Pi + T_j - \tau^h + x_j^G$ is computed as the sum of gross wages, pensions, capital income and net transfers T_j plus the sum of extracted local profits, that is, local property, land and business taxes. The term w_j is computed from income tax data IRCOM and national accounts as the sum of labor and pension income. Capital income Π is computed from national accounts as the sum of dividends, mixed income and rental income (including implicit rental income) net of total local household property tax computed from local tax data REI and BCC municipal accounts. It already accounts for national and local business taxes and is distributed lump-

²⁶In the last two terms in brackets, x^G , N and τ are bounded due to physical constraints. In most applications, agglomeration elasticity $1 + \epsilon_{G_k, N_k}$ will be bounded as well, and per capita expenditure x_j will be bounded from below by a strictly positive number.

sum in coherence with our hypothesis of geography-neutral ownership in Section 2.3.2 and the construction of composition-neutral variables.²⁷

Per capita net transfers from the central government T_j are computed from IRCOM and BCC data and include (i) taxes on wages, pensions and capital income (including means-tested negative income taxes), (ii) various annual operating subsidies to municipalities and municipal federations and (iii) the current *stock* of investment subsidies to municipalities and their federations discounted by $\frac{R}{1+R}$, the interest rate $R = 3\%$ hence transforming the current subsidy liability in a stream of yearly subsidies.

Municipal financial accounts allow us to construct x_j^G as the sum of operating expenditure and public assets, the latter being discounted by a factor accounting for depreciation and the fact that it is financed by debt repaid over time (see Jannin and Sotura 2019). We compute resident's tax share τ_j from REI local tax data and BCC financial account data as the sum of all taxes paid by residents — residence tax and resident user fees — divided by total municipal tax revenues — from residence, land, property, business and various resident and business user fees.

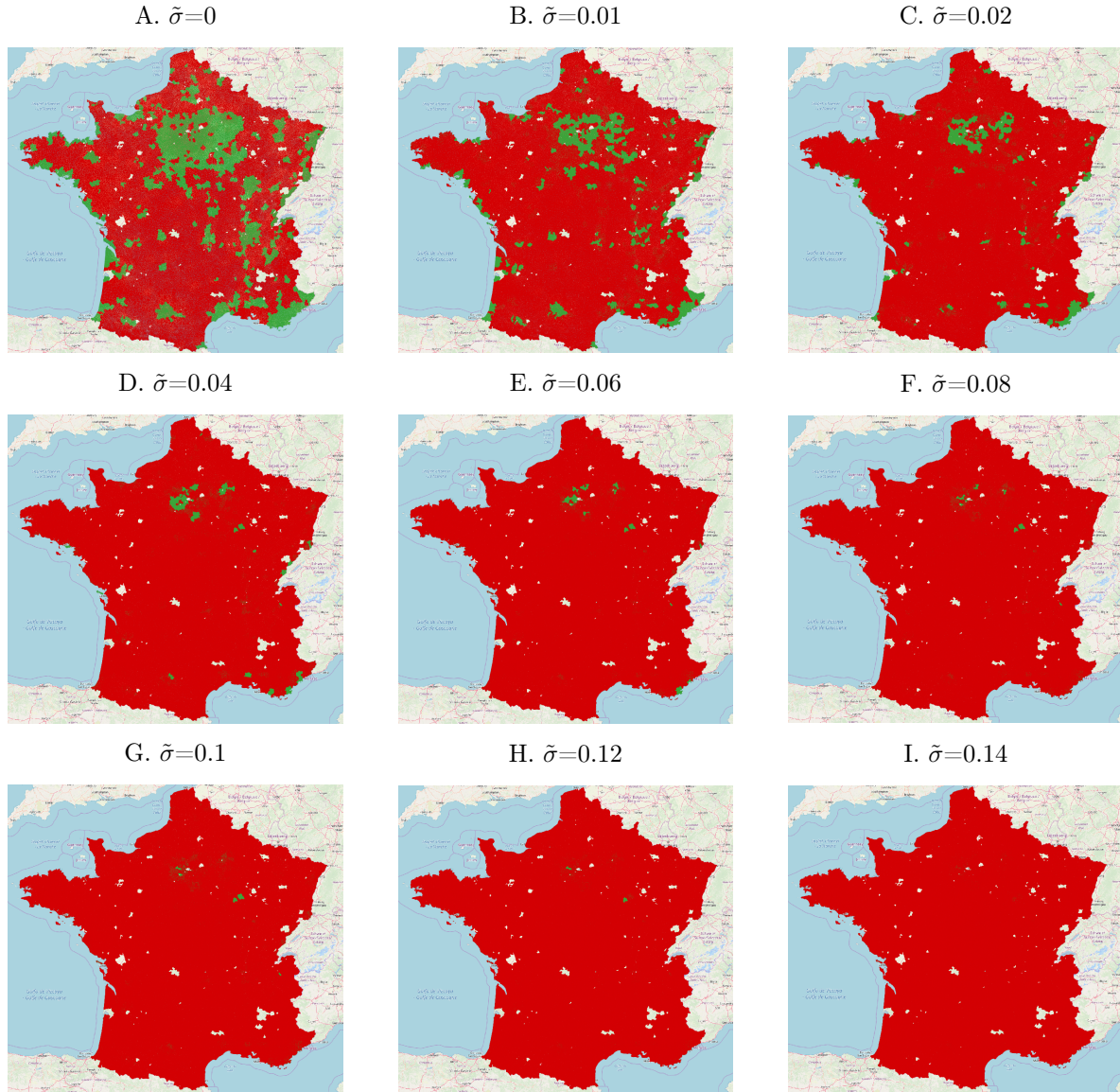
Estimate for scale economies To illustrate our main argument, we need an estimate for the agglomeration elasticity $1 + \epsilon_{G_j, N_j}$. For this empirical application, we assume that it is constant and equal to $1 - \kappa$ where κ is the congestion parameter. In their review, Ahlfeldt and Pietrostefani (2019) compute a public good density elasticity of 0.17 based on original estimates in Carruthers and Ulfarsson (2003). However, it is unclear whether this elasticity is comparable to what $1 - \kappa$ captures in our framework, which is the influence of city size on public good provision, holding constant urban land area. In appendix B.4, we calibrate our public good demand equation (2.21) in the spirit of the seminal work of Borchering and Deacon (1972) and Bergstrom and Goodman (1973) and report a value of $\kappa = 0.44$. We emphasize that this value is more reflective of association in the data rather than causal relationship, and use it to illustrate our main argument below.

The importance of preference heterogeneity We compute these implied weights pooling years 2010 and 2012 for different values of σ . Figure 2.5 offers compelling evidence that the set of Pareto-improving marginal reforms an economist would advocate strongly depends on

²⁷We make the admittedly strong assumption that VAT and other sales taxes are equivalent to a lump-sum tax on firms' profits and are already accounted for in our capital income term Π .

the amount of heterogeneity in location preferences $\tilde{\sigma}$ that she believes is responsible for the observed allocation.²⁸ Panel A maps the direction of marginal spatial transfers that a

Figure 2.5: Direction of Pareto-Improving Transfers ($\kappa = 0.44$)



Note: These maps show the sign of Pareto-improving marginal transfers that an economist would recommend conditional on different beliefs $\tilde{\sigma}$ for preference dispersion. Green areas should get positive marginal transfers and red areas should be marginally taxed (when all areas are red, no place should be taxed nor subsidized).

²⁸Note that for endogenous variables to stay at their observed level when σ changes, other deep fundamentals need to vary as well. However, the sufficient statistics approach precisely allows to characterize optimal policies independent of the underlying parameters, so that what parameters exactly are behind the observed equilibrium is irrelevant.

government who is blind to the existence of heterogeneity in location preferences would favor.²⁹ Regions in red would get less net transfers than in the current allocation, whereas regions in green would receive larger net transfers. Strikingly, under this homogeneous-preference hypothesis, marginal transfers should encourage migration to already dense and productive areas such as those surrounding France’s major cities like Paris, Lyon, Bordeaux, Strasbourg, Toulouse or Nice, as well regional subcenters such as Tours, Grenoble or Rennes. However as $\tilde{\sigma}$ increases, the range of marginal reforms that the government thinks are efficient narrows until it becomes empty, as exhibited in the other Panels of Figure 2.5. Hence, under-estimating preference heterogeneity may lead to advocating marginal reforms that would favour the most productive areas and would at best achieve redistribution, and at worse lead to efficiency losses.

To assess the efficiency of current transfers, one needs an educated guess for preference dispersion parameter σ . We take a central estimate from of Jannin and Sotura (2019) who find σ around 0.08 on similar data and at the municipal level. However, this estimate is valid only when locations are municipalities and not municipal federations. We show that under mild assumptions, we may still use it to carry out a back-of-the-envelope efficiency diagnosis.³⁰ Looking at Panel F in Figure 2.5, it is likely that the French current transfer system is close to being spatially efficient.

2.6 Equity and Density

We now turn to equity considerations and to the joint role of public good agglomeration economies and location preferences in defining the appropriate level of redistribution. In a context where high preference heterogeneity dampens migration responses, the set of transfers chosen by the government more likely reflects redistribution preferences than efficiency concerns. The intensity of observed transfers thus likely mirrors French preferences for spatial equity.

To provide intuition about how the interplay between public good agglomeration economies and location preferences may shape redistribution motives, consider the following two-city

²⁹This is obtained by letting $\tilde{\sigma} = 0$ which yields the efficiency condition $T_j - \sum_k N_k T_k = (1 - \kappa) (x_j^G - \sum_k N_k x_k^G) + x_j^G (1 - \tau_j) - \sum_k N_k x_k^G (1 - \tau_k)$.

³⁰Under the hypothesis that utilities are constant within municipal federations, the preference parameter at the municipal federation level σ_{mf} is close to the municipality-level parameter σ_m if the log difference between the number of municipalities in the federation N_{mf} and its geometric average across federations is small enough before the log difference between municipal federation population N_{mf} and its geometric average, so that $1 - \frac{\sigma_{mf}}{\sigma_m} = \ln\left(\frac{N_{mf}}{N_{mf}}\right) / \ln\left(\frac{N_{mf}}{N_{mf}}\right) \ll 1$, a condition that we empirically confirm.

example. Natural amenities in city 1 are exogenously better than in city 2, and it is also less costly to build housing there. Productivity is fixed, equal in both cities and production exhibits constant returns to scale so that wages are fixed and equal. Attractive and cheap city 1 will be more populated in equilibrium, and will offer more endogenous public good benefits because of positive agglomeration externalities. We can carefully engineer the fundamentals of the model—in particular the housing supply elasticity in 1 and 2—such that wages, rents, taxes and per capita consumption of housing and of the traded good will be exactly equal in equilibrium.³¹ The only difference between households will be the exogenous pure amenities and the endogenous public goods of the city they live in. In this context, one may want to compensate residents of 2 for benefiting from less public goods (and maybe for living in an unattractive place) because of horizontal equity concerns. Achieving this goal will require a tax and transfer scheme that compensates low density beyond income redistribution.

2.6.1 Compensation and Responsibility

Whether or not society wants to redistribute to low-income and low-density places will depend on the interpretation of idiosyncratic draws μ_{ij} determining agents' location in equilibrium. Are they spatial draws that should be compensated for—it may be costly to leave the environment one was socialized in—or pure preferences—one may really appreciate amenities in places destined to be low-density, e.g., remote steep mountains—that society deems the responsibility of individuals? The discussion here is reminiscent of “skill vs. laziness” type arguments found in the optimal taxation literature (see [Saez and Stantcheva 2016](#)) and more generally of the compensation vs. responsibility principle discussion (see [Fleurbaey and Maniquet 2011](#)).

Even if a burning issue for policymakers, the question whether preference draws pertain to the compensation or the responsibility principle is beyond the scope of this paper. What we propose to do however is to investigate what principle seems to empirically prevail as revealed by the current geographic tax and transfer scheme.

2.6.2 Revealed Social Preferences

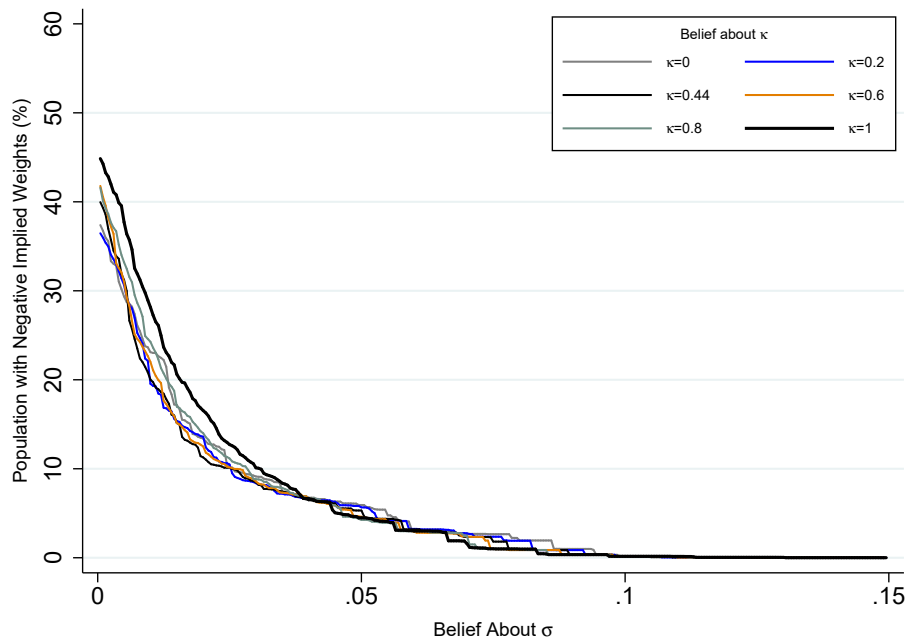
We use Proposition 8 to reveal collective preferences and investigate which underlying social welfare function is supported by the implied weights. However, when transfers achieve both

³¹For instance by making the extra assumption that amenities and public goods do not influence the marginal utility for consumption goods.

efficiency and equity objectives, one can only analyze the structure of implied weights *at an efficient allocation*. Indeed, since negative implied weights are a symptom of the government not accounting for all sources of inefficiencies when maximizing the social welfare function (2.29), they do not reflect actual social preferences.³² As such, any discussion on equity and revealed social social preferences should start by investigating the planner’s beliefs about sources of inefficiency, which must be such that the current situation is efficient from the point of view of a mistaken planner, i.e., is rationalized by positive implied weights.

Figure 2.6 reports the fraction of the population receiving negative weights as implied by the observed allocation, for different planner’s beliefs about preference dispersion $\tilde{\sigma}$ and congestion parameter $\tilde{\kappa}$. We report this relationship pooling years 2010 and 2012. Any com-

Figure 2.6: Government’s Beliefs



Note: This graph shows which government’s beliefs about σ and about κ are consistent with the observed transfers. We reject beliefs for which some implied weights are negative, as they would suggest that the planner is not acting rationally conditional on these beliefs.

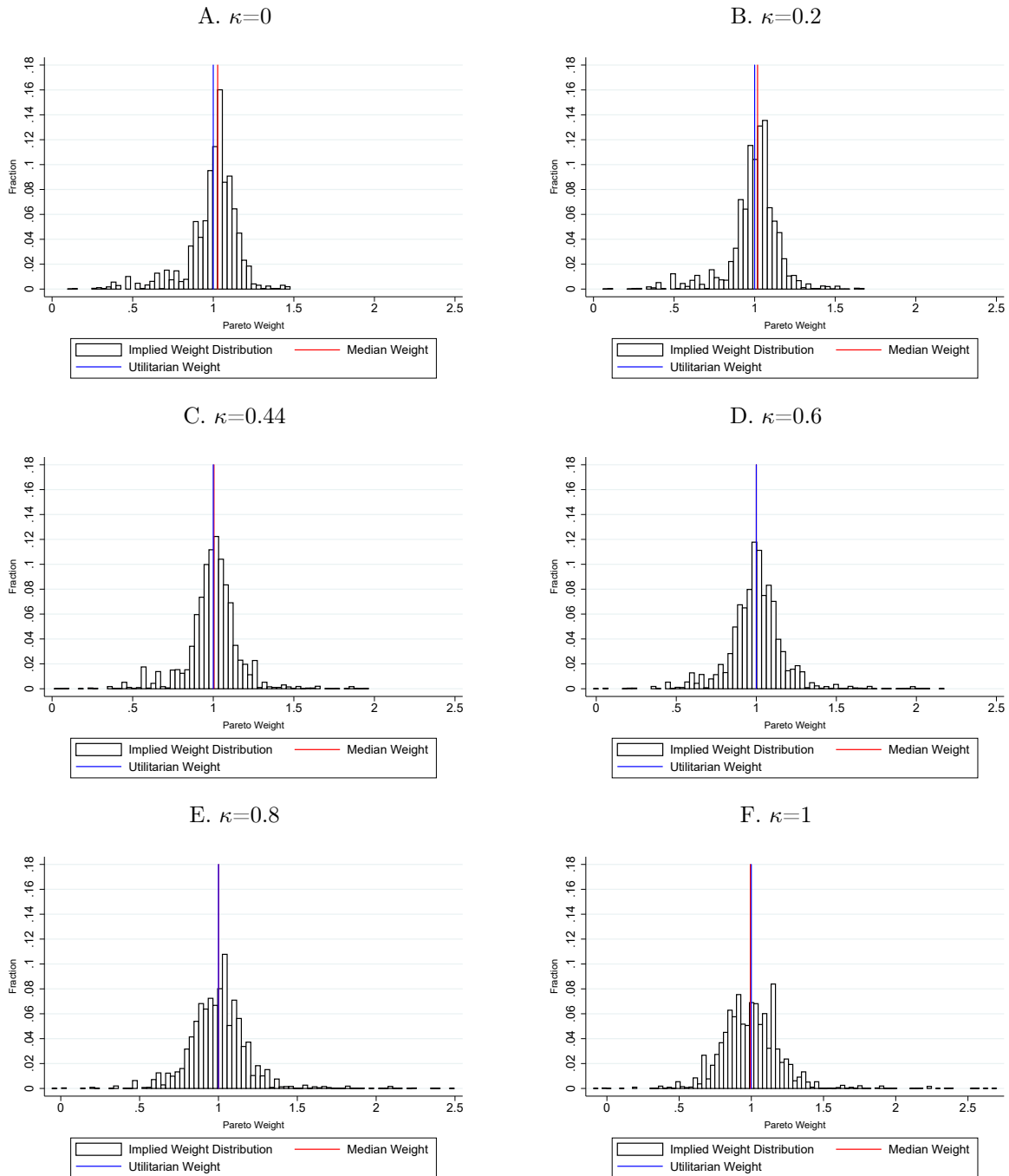
³²That is, if one accepts that the planner acts in accordance with the Pareto principle. Here we do not model how collective preferences emerge from the political process, however, assigning negative weights to a municipal federation would mean that the government wishes the welfare of individuals living in this jurisdiction to be less than what it could be, all others households’ welfare being held constant. Deliberate economic degradation of some locations does not seem to be consistent with any political process of preference aggregation, at least in a democratic system where there are no “second class citizens” nor spatial discrimination.

binations of $\tilde{\sigma}$ and of $\tilde{\kappa}$ such that implied weights are all positive may be what the government actually had in mind when designing “optimal” transfers leading to the observed situation. The only definitive conclusion one can draw from Figure 2.6 is that the government likely overstates preference dispersion compared to our best guess $\hat{\sigma} = 0.08$, as only values for $\tilde{\sigma}$ above 0.13 rationalize all possible planner’s optimizations.

We make the conservative assumption that the planner is as close as possible to our best guess $\hat{\sigma} = 0.08$ while being rational, i.e., we assume that the planner has belief $\tilde{\sigma} = 0.13$. It seems reasonable to assume that the democratic process of preference aggregation leads to the government having beliefs about σ not too far from the “true” value since location preferences are a common feature of all individuals. For $\tilde{\sigma} = 0.13$ we cannot reject any belief $\tilde{\kappa}$ since they all lead to positive implied weights. We then have to examine their structure for each situation. Figure 2.7 reports the distribution of individual Pareto weights for various beliefs about the true parameter κ . Strikingly, for most beliefs the median of the weight distribution is almost exactly equal to one, i.e., to the utilitarian weight, and the weight distribution is quite symmetric. For all beliefs, the dispersion around the median is quite limited with 90% of the population being weighted within 30% of the unitary weight. It thus appears that the French government is roughly utilitarian and that it carries out some limited redistribution beyond what a strictly utilitarian objective would command.

We investigate the structure of this deviation from utilitarianism. We test two simple hypothesis: whether the government compensates low resources beyond utilitarianism—which already seeks to compensate high marginal utility of consumption—and whether it compensates low density which we believe captures limited access to local public goods. Pooling years 2010 and 2012, Table 2.1 reports results of the corresponding regressions. We cannot conclude that the planner compensates low resources beyond the utilitarian objective. However, whatever the governments’ true belief about the congestion parameter, we estimate that it compensates lower densities, i.e., lower population sizes holding constant urban sprawl, and larger sprawl holding constant population size. Holding constant resources and urban sprawl, a doubling of status quo population size increases Pareto weights by 2% to 8%. This last piece of evidence supports some kind of horizontal equity concerns for the consumption of local public goods. We argue that public good agglomeration economies likely explain the horizontal inequities that the government seems to compensate here.

Figure 2.7: Distribution of Implied Weights for Various Planner's Beliefs



Note: These graphs show the distribution of the planner's individual Pareto weights implied by the observed 2010 and 2012 allocations under different planner's belief for congestion parameter κ at $\bar{\sigma} = 0.13$. The distribution of individual weights is obtained by weighting MF Pareto weights by MF population.

Table 2.1: Structure of Pareto Weights

	$\tilde{\kappa} = 0$	$\tilde{\kappa} = 0.2$	$\tilde{\kappa} = 0.44$
Log Resources	-0.235 (0.288)	-0.146 (0.261)	-0.110 (0.273)
Log Population	-0.027** (0.014)	-0.042** (0.016)	-0.055** (0.023)
Log Urbanized Land	0.012 (0.015)	0.034** (0.016)	0.057*** (0.022)
R-squared	0.055	0.057	0.053
Observations	4, 133	4, 133	4, 133
	$\tilde{\kappa} = 0.6$	$\tilde{\kappa} = 0.8$	$\tilde{\kappa} = 1$
Log Resources	-0.092 (0.283)	-0.072 (0.297)	-0.051 (0.317)
Log Population	-0.062** (0.028)	-0.071** (0.033)	-0.080** (0.040)
Log Urbanized Land	0.071*** (0.026)	0.089*** (0.032)	0.109*** (0.038)
R-squared	0.050	0.046	0.041
Observations	4, 133	4, 133	4, 133

Note: This Table explores the structure of Pareto weights for different hypothesis about the planner’s belief for congestion parameter κ . Per capita resources are constructed as in Section 2.3.

2.7 Conclusion

In this paper, we study the implications of heterogeneous location preferences for the design of efficient and equitable spatial transfers when local public goods are the leading cause of migration externalities.

Using a unique combination of French local administrative panel datasets, notably comprehensive local financial accounts, we first document new stylized facts suggesting the existence of large public good scale economies.

We then build a spatial equilibrium model that incorporates the most salient features of the local tax system in France. We characterize the transfers that implement efficient population distributions by the way of estimable sufficient statistics. We show that with homogeneous preferences, the laissez-faire economy is generally inefficient and there is room for welfare-enhancing spatial transfers. However, Pareto-improving transfers may not exist when location preferences are heterogeneous, because heterogeneity increases the equity cost of migration (transfers) relatively to its efficiency benefits (agglomeration gains). This has

important consequences for the design of spatial policies aiming at harnessing agglomeration forces for the benefits of all.

We then apply our framework to the current French transfer system. We empirically show the importance of accounting for location preferences as ignoring them may lead to severely mistaken recommendations that overstate the feasibility of marginally efficient reforms and to anti-redistributive policies. Finally, we explore whether the coexistence of location preferences and public good agglomeration economies creates horizontal equity concerns about public good consumption and find that the French system indeed compensates low-density areas beyond a utilitarian objective.

Overall, our research suggests the need for more work on the estimation of public good scale economies — and more broadly on the impact of density on consumption amenities — and on the determinants of location preferences, as their interaction bears important consequences for the design of efficient and equitable spatial policies.

Chapter 3

The Deadweight Loss of Property Transaction Taxes

This paper studies the impact of transaction taxes in the property market. I use a unique administrative dataset covering all property transactions in France over the period 2010–2016 and exploit quasi-experimental variations provided by a 2014 reform that increased stamp duties. I first document that tax changes were salient and led to significant re-timing responses. Exploiting spatial differences in the timing of tax increases, I then use a synthetic control method to investigate extensive responses and estimate a net tax elasticity in the range $[0.22, 0.25]$. I find no effect on transaction prices nor house quality, suggesting that the supply of existing homes may be very elastic. I propose an alternative mechanism in a search model in which tax increases may upgrade average match surplus and push bargained prices up by discouraging buyers with low reservation values. This improved matching efficiency may positively contribute to welfare.

3.1 Introduction

Property transaction taxes exist in many developed countries and are long-rooted in their tax history. In the UK for instance, fixed stamp duties on parchment and vellum were introduced in 1694 and gradually extended to various consumption goods, until they started being levied on property transfers in 1808. In France, the government imposed a flat 4% rate on the value of property transactions in 1790 during the Revolution, thereby unifying a myriad of existing small taxes (Friggit 2003). It later reached record heights during the period 1920–2000, peaking at more than 35%. While fiscal capacity constraints certainly explain the prevalence of transaction taxes in the past, their continued existence today likely results from the interplay between political economy—local authorities wish to retain tax autonomy—and local tax systems’ path-dependence to sunk administrative costs. Nowadays, property transaction taxes represent an important source of funding for French local jurisdictions and contributed EUR 11 billion (0.5% of GDP) to the financing of various local public goods in 2016.

Although some have recently argued for Tobin taxes in order to curb speculation in asset markets (e.g., Dávila 2014), economists traditionally oppose property transaction taxes as they introduce obvious distortions. Indeed, they create straightforward welfare losses from asset misallocation as mutually beneficial exchanges are kept at bay. Transaction taxes are also suspected to create lock-in effects dampening residential mobility. They entail geographic misallocation of households and may adversely affect labor market outcomes by increasing spatial mismatch, although this latter inefficiency is likely of second order given the large gains from long-distance labor-related moves (Hilber and Lyytikäinen 2017).

In this paper, I provide quasi-experimental evidence on behavioral responses to property transaction taxes. In 2014, a reform let French counties (*départements*) increase stamp duties (*droits de mutation à titre onéreux* or DMTO) levied on property transactions from 5.09% to 5.81% of the purchase price. Initially envisioned as a temporary budgetary boost, tax increases were made perennial shortly after the first implementations. I show that transaction taxes are distortive across a range of margins by studying bunching and extensive responses to tax changes.

Other papers have investigated the French 2014 stamp duty reform (Bachelet and Poulhès 2017; Bérard and Trannoy 2018). I improve on this research in two ways. First, I obtained first-time access to a unique administrative dataset recording all property transactions in

France over the period 2010–2016 (*Demandes de Valeurs Foncières* or DV3F) featuring exact sale date, selling price and property characteristics. Coverage is exhaustive because property transactions are subject to taxes collected by the government on behalf of local authorities. Any buyer-seller pair must sign the final deed of sale under the supervision of a publicly appointed notary who then passes transaction terms on to the government. Second, I exploit the geographic granularity of the data to carry out a synthetic control estimation of extensive responses that exploits differences in the timing of reform adoption across mainland French counties.

When voted for by a county council, the tax increase was announced between one and two months prior to its effective implementation date thereby creating a time notch below which buyers and sellers were incentivized to transact. I first carry out a bunching exercise to investigate short-term re-timing responses to the tax increase. I find that the 0.72 percentage points (pp) tax increase was salient and anticipated. It led transaction partakers to close deals an average 18 days earlier and up to 57 days. Re-timing responses increased market activity in the average county by 76% in the week prior to the reform and depressed it by 31% in the following two months relative to a counterfactual post-reform period with extensive responses only. I then investigate the presence of spatial bunching, i.e., the existence of excess and missing transactions around notches created by static transaction tax wedges at treatment administrative borders. I find no evidence of such re-location spillovers, suggesting that tax avoidance responses were mostly re-timing responses.

Exploiting the phasing-in of the reform's adoption across French counties, I implement a synthetic control method on low-level geographic treatment and control units by aggregating transaction data at the municipal federation (MF) level. These geographic units are spatially nested in larger treatment groups that encompass several French counties forming large property markets operating under a common tax regime. I first present difference-in-differences (DiD) type evidence comparing average treated and synthetic transaction trajectories in levels. I then turn to the cumulative transaction gap between average treated and average synthetic units which neutralizes re-timing responses and reveals the exact amount of extensive responses. I precisely estimate that the 0.72pp tax increase depressed market activity by an average 3.1%–3.5%, which corresponds to a net tax elasticity in the range $[0.22, 0.25]$. I find no effect on average prices nor on average quality of traded homes.

A standard explanation for this price result is that supply of existing homes is very elastic,

even in the short run. I illustrate an alternative mechanism in a stylized matching framework where buyers have heterogeneous reservation values. A transaction tax increase may screen low-value buyers out of the market, thereby pushing up bargained prices through a composition effect. I discuss welfare implications in this simple setting. While increasing taxes may decrease the number of mutually beneficial transactions and lead to efficiency losses, it may improve allocative efficiency by increasing gains from trade through changes in the composition of active buyers. In this simple setup, I find that this positive composition effect likely outweighed the negative lock-in effect following the 2014 reform. I emphasize that this result obtains under deliberately simplifying assumptions and mainly illustrates the importance of accounting for buyer heterogeneity in matching markets such as the housing market.

I contribute to the small but growing empirical literature studying behavioral responses to property transaction taxes which I briefly review here. While optimization responses (bunching around time and price notches) are pervasive, the extent of “real” extensive responses and price capitalization very much vary across housing markets and research designs. [Benjamin et al. \(1993\)](#) estimate that the 1.6pp transfer tax increase (from 3.5% to 5.1%) in 1988 Philadelphia fully capitalized into prices. [Van Ommeren and Van Leuvensteijn \(2005\)](#) find that a 1pp increase in transaction costs is associated with an 8% decrease in residential mobility to ownership in the Netherlands. [Dachis et al. \(2011\)](#) implement a DiD strategy on a border discontinuity to study the Toronto Land Transfer Tax. The authors find that the newly introduced 1.1% average transfer tax on top of the existing 1.1% average provincial tax depressed transaction volume by 15% and that its incidence entirely fell on sellers. Considering other transaction-related costs, they estimate an elasticity of transactions to transaction costs of about 1. [Davidoff and Leigh \(2013\)](#) study stamp duties in Australia and find that it impacts extensive margin choices as well as housing prices, with incidence entirely falling on sellers.

[Besley et al. \(2014\)](#) study the incidence of the UK 2008–2009 stamp duty holiday in a structural bargaining framework and find that the 60% of the tax cut accrued to buyers. The authors find significant evidence of transaction re-timing before the holiday ended but not evidence of extensive margin responses. [Best and Kleven \(2017\)](#) examine the same reform but do estimate extensive responses. Similarly to [Besley et al. \(2014\)](#), they implement a difference-in-differences strategy using price groups and find that the 1% tax cut stimulated market activity by 20% during the holiday period, more than half of which came from extensive margin responses.

As mentioned, two studies are the closest to this paper. [Bérard and Trannoy \(2018\)](#) study the 2014 French stamp duty reform using county-level data on monthly transaction tax bases. They estimate significant re-timing responses, but find no evidence of medium-term extensive responses. [Bachelet and Poulhès \(2017\)](#) also investigate the 2014 reform, gathering data from French notary offices. However, the authors identify an average 10% medium-term decrease in transaction volumes across municipalities following the tax increase, and that tax incidence entirely fell on buyers.

I also relate to the literature on bunching (see [Saez 2010](#); [Chetty et al. 2011](#); [Kleven 2016](#)). Some studies explicitly investigate bunching behaviors around time or price notches. [Slemrod et al. \(2017\)](#) estimate significant transaction re-timing and price bunching around the introduction of a 0.8pp tax notch at a reference price in Washington D.C., but find no extensive margin response. [Kopczuk and Munroe \(2015\)](#) also estimate significant price bunching around a notched transaction tax in New York and New Jersey, with the tax burden falling on sellers.

This paper contributes to the literature on synthetic control methods for regional policy evaluation pioneered by [Abadie and Gardeazabal \(2003\)](#) and further developed in [Abadie et al. \(2010\)](#) and [Abadie et al. \(2015\)](#). I notably build on methodological contributions and applications that generalize the method to multiple case studies (e.g., [Cavallo et al. 2013](#); [Gobillon and Magnac 2016](#); [Acemoglu et al. 2016](#)).

Finally, this paper pertains to the literature on allocative efficiency in the housing market (e.g., [Glaeser and Luttmer 2003](#)) and on the role of taxation and policy instruments more broadly in matching markets ([Lundborg and Skedinger 1999](#); [Shimer and Smith 2001](#); [Landais et al. 2018](#); [He and Magnac 2018](#)).

The rest of the paper proceeds as follows. Section 3.2 gives background on the French stamp duty and presents the data. Section 3.3 examines bunching responses. Section 3.4 investigates extensive responses. Section 3.5 discusses welfare implications in a search model. Section 3.6 concludes.

3.2 Institutional Background and Data

3.2.1 French Administrative Geography

The French territory breaks down into four spatially nested layers of local governance. The 35,000 *communes* (henceforth municipalities) are the smallest administrative jurisdictions, nested in 2,000 *intercommunalités* (henceforth municipal federations or MFs) envisioned as an

intermediate step between reinforced municipal cooperation and complete merger. A *département* (henceforth county) is itself a collection of municipal federations (96 counties excluding overseas territories), and a *région* (13 excluding overseas) a collection of counties. Each local authority has some autonomy to raise local taxes. For instance, municipalities may raise taxes on land, residence, property and businesses. In particular, counties have the power to levy taxes on property transactions, i.e., stamp duties.

3.2.2 The French Stamp Duty

The *droits de mutation à titre onéreux* (DMTO) is the French stamp duty that applies to the total value of any built property transaction in the market for existing properties. Its statutory incidence falls on buyer. The stamp duty is managed and collected by the Ministry of Finance through notary offices on behalf of local authorities and paid back to the latter. For transaction partakers, it is compulsory to sign the deed of sale under the supervision of a notary who guarantees its authenticity and handles its legal aspects. Information on each deed of sale is then passed on to the governmental Land Registration Service. Total tax liability is obtained by multiplying the transaction price by a flat rate which is—up to December 31st 2013—the sum of

- (i) A fixed municipal rate of 1.2%. This rate is not subject to modification by the municipal council. Tax proceeds are allocated to the municipality in which the transaction takes place if it has more than 5,000 inhabitants, and to a county equalization fund if not.¹
- (ii) A discretionary county rate anywhere between 1.2% and 3.8%. The county rate is decided by the elected county council and is free to change year-on-year. In practice, all counties had been applying the maximum admissible rate of 3.8% between January 1st 2010 (at least) and December 31st 2013. Tax proceeds are allocated to the county the transaction takes place in.
- (iii) A national rate which is 2.37% of the county rate ie $2.37\% \times 3.8\% = 0.09006\%$. Tax revenues are allocated to the government to cover the administrative costs of collecting taxes on behalf of municipalities and counties.

The comprehensive rate that applies to standard built property transactions in the secondary market (*taux de droit commun*) as of December 31st 2013 is then 5.09006%.² Prior to this

¹*Fonds départemental de péréquation.*

²Land and newly built properties are subject to a smaller 0.715% stamp duty and to a 20% VAT rate.

date, stamp duties had been constant since January 1st 2006 when the government 2.37% rate was introduced (Friggit 2003). In 2013, total DMTO proceeds were EUR 8.5 billion (0.3% of GDP).³ A buyer of the 2013 average residential property faced a tax liability of EUR 9,560.⁴

3.2.3 The 2014 Stamp Duty Reform

On December 29th 2013, the French parliament voted the 2014 national budget which was published in the Official Bulletin on December 30th 2013. The 2014 budget notably raised the maximum admissible rate for the DMTO county component, from the initial rate of 3.8% to a new maximum rate of 4.5%. The reform did not impose a tax increase as it did not change the discretionary power of counties over county rate. County councils were let free to vote for a new county rate up to the 4.5% limit. The only other limitation was that any new rate would be enforced no sooner than the first day of the second month following the county council vote. The earliest possible council meeting after the Budget was passed being January 2nd 2014, the earliest date at which property transactions could be taxed at a modified rate was March 1st 2014.

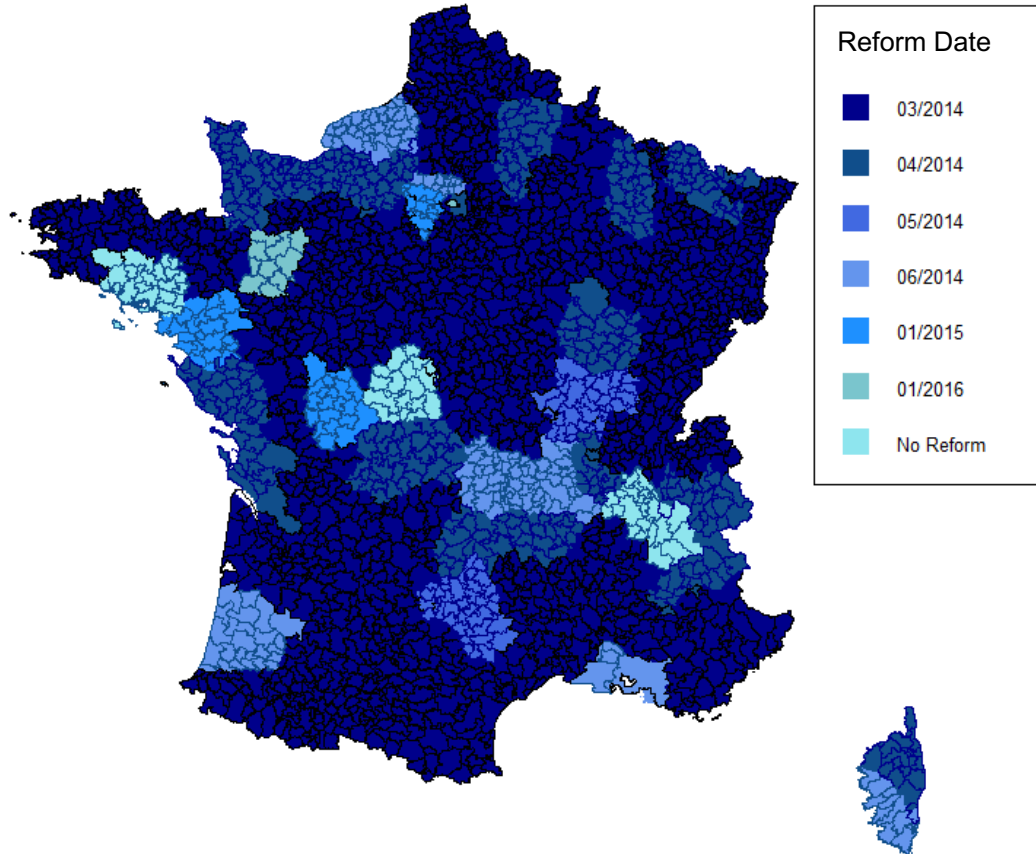
In practice, all counties that decided to raise their rate applied the new 4.5% maximum rate. In counties that did so, the national rate was mechanically modified following the change in the county rate from $2.37\% \times 3.8\%$ before the reform to $2.37\% \times 4.5\%$ after. Overall, transactions taking place in counties that implemented the reform saw proportional taxation increase from 5.09006% to 5.80665% of the transaction price. In 2016, total DMTO proceeds were EUR 11 billion (0.5% of GDP). The average 2016 residential property transaction — if taxed at the new rate — faced a tax liability of EUR 10,960.

Not all counties raised stamp duties on March 1st 2014. Reform adoption was phased-in over 2014, 2015 and 2016 with different groups of counties adopting the new 4.5% rate at different dates. I henceforth interchangeably use “county group” or “treatment group” to designate one of the 7 groups of counties defined by their treatment date, group number 7 being the group of untreated counties as of January 2017. Figure 3.1 shows the geography of county groups as well as the administrative boundaries of their member municipal federations. Groups {1, 2, 3, 4, 5, 6, 7} respectively represent {60; 19; 2; 7; 3; 2; 3} counties and {1, 324; 408; 68; 151; 56; 12; 66} municipal federations. Figure 3.2 shows the time-line of reform adoption across groups. Reform implementation date was respectively March 1st, April 1st,

³Source: French Municipal Financial Accounts.

⁴Source: see data section below.

Figure 3.1: Treatment Groups

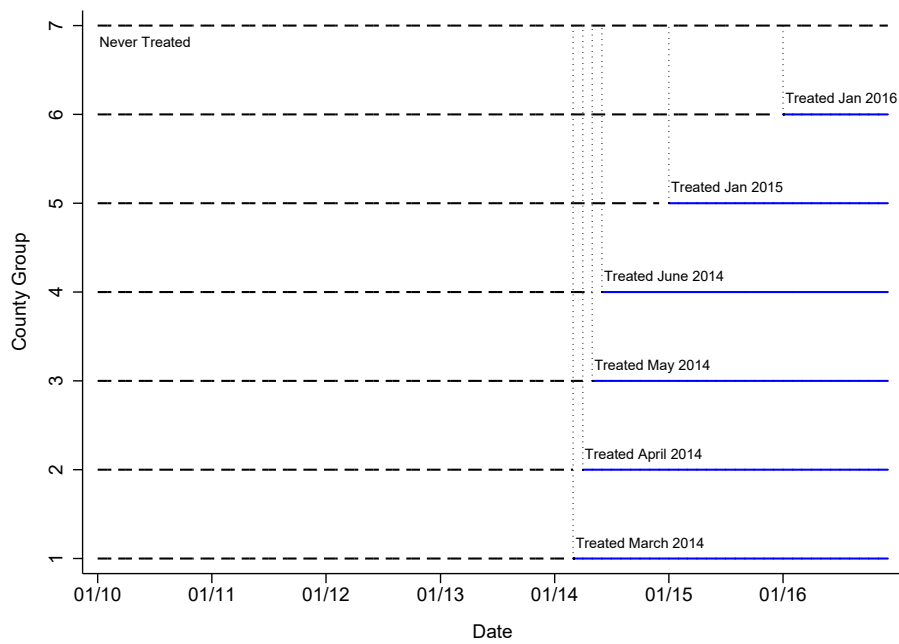


Note: This Figure shows the geography of county treatment groups. I report in solid lines the administrative boundaries of the 2,000 municipal federations in 2014.

May 1st and June 1st 2014 for groups 1, 2, 3 and 4. It was January 1st 2015 for group 5 and January 1st 2016 for group 6.

The 2014 budget law stated that any tax increase implemented after March 1st 2014 would be fully reverted on March 1st 2016. Indeed, the reform was introduced to temporarily help fund counties' post-crisis budget unbalances. However, the 2015 budget law voted on December 29th 2014 confirmed shortly after that any tax increase implemented before March 1st 2016, the initial end date of the tax increase window, would be made perennial.

Figure 3.2: Reform Adoption Timeline



Note: This Figure shows the phasing-in of the 2014 DMTO reform across county groups over time, group 7 being never treated over the period 2010–2016.

3.2.4 Data and Descriptive Evidence

Datasets I obtained first-time access to a unique administrative property transaction dataset obtained from the French Ministry for the Environment called *Demandes de Valeurs Foncières* (DV3F) for the period 2010–2016. As mentioned, property transactions are subject to taxes collected by the central government through notary offices on behalf of local jurisdictions. Furthermore, property ownership changes must be reported to the central government who also manages the local property tax register. For these reasons, DV3F provides an exhaustive coverage of all transactions on the primary and secondary French property markets, that is, all land and built property transactions involving natural persons, businesses or public entities.

DV3F features detailed information on each transaction including its price, the property’s surface and many other quality indicators. For residential estates, it reports for example the number of living rooms, bedrooms, garages, terrace or swimming pools. For pure business estates, additional quality is mostly the existence of a garage. Importantly, the dataset provides transaction spatial details at an infra-municipal level. There is however a geographic and

historical exception: DV3F does not cover the 3 border counties that were lost to Germany during the 1870 Franco-Prussian war and reintegrated in 1919 after WWI.⁵

I supplement DV3F transaction data with administrative tax data giving information on average wage and pension income at the municipal level for all French municipalities and all years from 2004 to 2015 (*Impôt sur le Revenu par Commune* henceforth IRCOM). I also use comprehensive data on municipal population and housing stock (*Fichier des Logements à la Commune* henceforth FILOCOM) covering all municipalities every even year from 2002 to 2014.⁶ I use FILOCOM to construct various municipality-level variables on population and housing stock. I finally use publicly available data on the geographic and administrative partition of mainland France, as well as socio-demographic and census data from the French National Statistical Institute (INSEE) for 2008.

Sample selection Regarding DV3F, I focus on transactions that are taxed under the standard regime (*régime de droit commun*), i.e., built property transactions concerned by the 2014 DMTO reform.⁷ They correspond to standard private transactions in the secondary residential property market—flats and houses—as well as a minority of business estate transactions. I further drop the latter to focus on residential estates only. This leaves me with around 4.4 million transactions over the period 2010–2016.

As for geography, I restrict my analysis to transactions occurring in mainland municipalities that experience no boundary changes between 2010 and 2016 included. There are 36,682 French municipalities in 2010, of which 34,576 are on the mainland and covered by DV3F (i.e., after the exclusion of Corsica, overseas territories and the 3 counties bordering Germany). Finally, 34,378 municipalities do not experience boundary re-definitions (mergers, split-ups, etc.) between 2010 and 2016 included. This last conservative restriction will be useful for the aggregation of transaction-level information at the municipal federation level in Section 3.4.

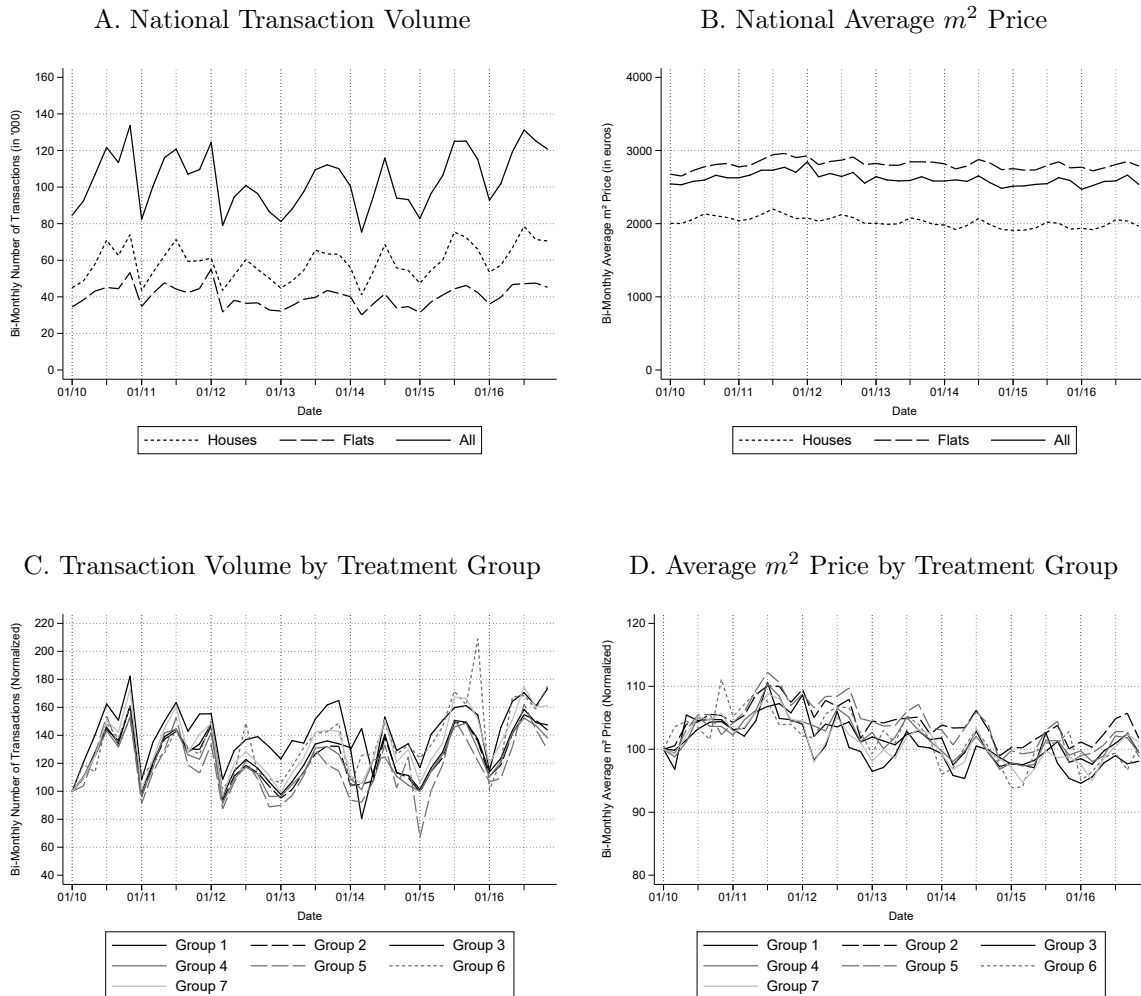
Raw evidence Panels A and B of Figure 3.3 visually inspect the raw number of transactions and average m^2 price at the national level for my final sample of transactions and municipalities, broken down by dwelling type (flat or house). They offer evidence that the housing

⁵These counties are *Moselle*, *Bas-Rhin* and *Haut-Rhin*. They are subject to derogatory local laws, in particular with regards to the local property register and the collection of transaction taxes. Transaction data for these counties are available for consultation only locally.

⁶Information on dwelling characteristics present in housing stock data FILOCOM and transaction data DV3F comes from a same source database: the national property register, managed by the government.

⁷Corresponding to tax code (*article CGI*) 1594D*1.

Figure 3.3: Summary Statistics



Note: This Figure reports volume and m^2 prices raw series. The Paris municipality is excluded in all four Panels. Panels A and B show total transaction volume and average m^2 price at the national level, for all residential dwellings and by dwelling type. Average m^2 price is computed as the two-month sum of m^2 prices across all transactions divided by the two-month total number of transactions. Panels C and D show the two-month total number of transactions and average m^2 price by treatment group, normalized with respect to January–February 2010.

market did not experience significant shocks between 2010 and 2016, barring seasonal cycles.

Two nation-wide reforms implemented during this period are worth mentioning. The first raised the baseline tax rate on property capital gains from 28.1% to 31.3% — except for main residences, for which capital gains were and stayed fully tax exempt — and was applicable to sales happening from January 1st 2011 onward. The second, implemented in February 1st 2012, extended the detention period over which property capital gains gradually become tax

exempt from 15 to 30 years, again without changes for main residence capital gains. The reform applied retro-actively to all properties and was anticipated so that households who had owned property for more than 15 years faced a strong incentive to sell before February 1st 2012. These two reforms likely explain the most salient spikes in Panel A which are presumably anticipation responses.

Panels C and D show volume and m^2 price trends by treatment group pooling all dwelling types. All seven treatment groups faced a similar evolution over the period 2010–2016 in terms of both volume and prices. In all four Panels, I exclude the Paris municipality which is an outlier regarding m^2 price trend. Together, these raw trends offer reassurance that the period 2010–2016 is not the scene of significant turmoils in the housing market and is well-suited for investigating the effect of a modest transaction tax reform.

3.3 Bunching

I expect a stamp duty increase to prompt three kinds of behavioral responses. First, implementation dates create time notches below which buyers and sellers are induced to re-time transactions. Second, spatial differences in the timing of reform adoption create geographic tax wedges. These distortions will encourage some buyers—those initially marginally indifferent between two (or more) locations plus some beyond—to bunch at the treatment border as they will try to escape the new tax while buying as close as possible to their preferred residential or investment location. Third, one expects to see extensive margin responses (the standard lock-in effect) as some transactions that would have occurred absent the tax increase do not take place because of it. In this section, I investigate the intensity of re-timing responses and border bunching.

3.3.1 Bunching at the Time Notch

In France, the period between a buyer-seller formal match—the signature of the preliminary sale agreement—and the actual transaction date—the date the bill of sale is signed, which is relevant for tax purposes—is negotiated between transaction partakers. It gives buyers time to contract a mortgage and potentially withdraw, and lasts on average 3–4 months including a 10-day minimum legal period. Assume for exposition that the 3-month period holds exactly for all potential buyers and sellers. The day the reform is voted by the county council, we can assign county residents to two possible types. Households not searching on the market—

because they thought of searching later, including never—or already searching but not yet matched with a seller have a maximum of two months—this is if the county council’s vote takes place on the first day of the month—to find a transaction partner and sign a bill of sale if they want to avoid the tax increase, which is impossible. Hence, only residents already in a buyer-seller match are able to move the transaction date forward by acting diligently regarding paperwork.

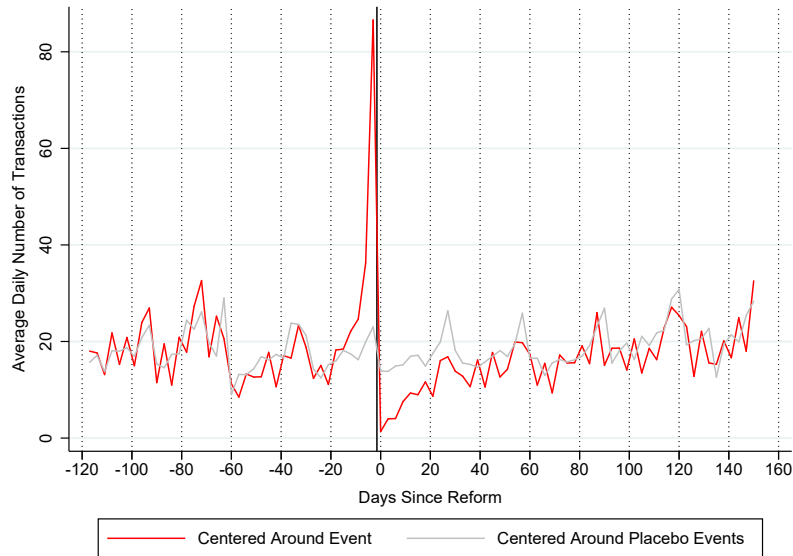
Six reforms taking place at six different dates in six different treatment groups can be superimposed to estimate the average timing response across treated counties. The small magnitude of the reform suggests modest behavioral responses in absolute value, which imposes to build a distribution that gets as close as possible to the average daily transaction count to avoid smoothing away timing responses.

For the bunching exercise, I aggregate the daily transaction count at the county level. For each county-date observation, I first define the distance d (in days) to its actual reform date. Distance belongs to a 363 day window around the reform, i.e., $d \in [-150, 212]$, is negative for counties observed before their reform date and positive for counties observed after. I similarly construct distances $d \in [-150, 212]$ to placebo reform dates, defined for each county as the day and month of its actual reform date in a different year. A county is hence observed a least five times at any $d \in [-150, 212]$ in my dataset, once for the distance to its actual reform date, and four times for distances to placebo reform dates. I finally restrict the sample to form a balanced panel of county-distance observations, that is, the set of observed counties is the same at each d across (placebo) reforms.

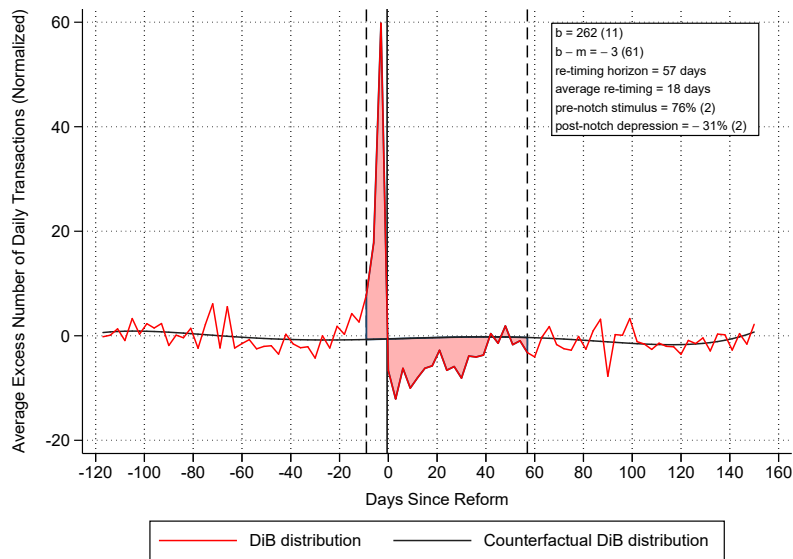
A naive approach consists in superimposing the six reforms by computing an average transaction count by distance bin across the different counties and reforms. For smoothness purposes, I group distances in bins of 3 days: bin $j = 0$ includes the day of the reform and the following 2 days, bin $j = 3$ the fourth day and the following 2 days, etc. I restrict the graph to $j \in [-120, 150]$ and bin together all observations below $j = 120$ in bin $j = 120$ and above $j = 150$ in bin $j = 150$. Panel A of Figure 3.4 shows in red the average county daily transaction count in each distance bin, pooling the six reforms. In grey is the average daily transaction count taken around placebo reform dates across treated counties. As is standard, top and bottom distance bins are dropped. The number of transactions surges in the days before the reform is implemented while there is a much weaker similar pattern around placebo reform dates.

Figure 3.4: Bunching at the Time Notch

A. Naive Bunching



B. DiB Estimates



Note: Panel A shows in red the distribution of the daily transaction count around reform date averaged across treated counties, pooling all events across treated groups. In light grey I report the raw transaction distribution averaged across windows around placebo reform dates. Panel B shows DiB estimates for the intensity of transaction re-timing around reform dates, pooling all reforms across treated counties. Standard errors are computed by bootstrapping residuals 1,000 times.

This simple approach likely leads to biased bunching estimates since daily transaction count is affected not only by the tax reform, but also by potential day-of-the-week, day-of-the-month, and bank holidays shocks.⁸ More generally, cycles in the housing market are likely to bias naive bunching estimates. To deal more generally with these issues, I construct a distribution of daily transaction count which is stripped of these potential sources of biases by implementing a flexible difference-in-bunching (DiB) strategy. I run the following regression pooling all county-date observations:

$$\begin{aligned}
 n_{it} = & \sum_j \beta_j^T \mathbf{I}_{d_{it}=j} \times \text{Treat}_{it} + \sum_j \beta_j \mathbf{I}_{d_{it}=j} + \eta \text{Treat}_{it} + \sum_{k=1}^7 \gamma_k \mathbf{I}_{\text{dow}_{it}=k} \\
 & + \sum_{l=1}^{31} \delta_l \mathbf{I}_{\text{dom}_{it}=l} + \sum_{m=1}^{12} \theta_m \mathbf{I}_{\text{month}_{it}=m} + \mu \text{Holiday}_{it} + \alpha_i + \varepsilon_{it}
 \end{aligned} \tag{3.1}$$

where n_{it} is the number of transactions in county i at date t and d_{it} is the distance to the reform date or to the placebo reform date. Treat_{it} is a dummy for the county-date observation being in the $d \in [-150, 212]$ window around its actual reform date. Treat_{it} allows me to carry out a DiB estimation by comparing market activity between treated and control distance windows. The remaining terms are a set of day-of-the-week, day-of-the-month, month, bank holiday and county fixed effects. This specification uses only the treatment window when estimating the excess number of transactions around the reform date, and all county-date observations to smooth cycles away. For the average county, $\hat{\beta}_j^T$ hence represents the estimated impact of being j days from its actual reform date on its daily transaction count, relatively to what's typically happening j days from placebo reforms and stripped of housing market cyclical effects and other time and county fixed effects.

Although the empirical distribution is affected by extensive margins responses, I only need a partial counterfactual to estimate the intensity of timing responses, i.e., a counterfactual stripped of all intensive timing responses but not of extensive responses (Kleven 2016). I fit a flexible polynomial on the unbiased empirical distribution, excluding a region before and after the notch. The excluded region's lower and upper bounds are chosen to minimize the difference between bunching and missing transaction masses around the time notch. The identification assumption is the smoothness of the counterfactual at the notch stemming from the absence

⁸Two of the four reforms take place on weekend days, and one takes place on a bank holiday.

of extensive margin responses in its vicinity. Formally, I run

$$\hat{\beta}_j^T = \sum_{k=0}^6 \beta_k(j)^k + \sum_{l=\underline{d}}^{\bar{d}} \gamma_l \mathbf{I}_{d_j=l} + \varepsilon_j \quad (3.2)$$

where $\hat{\beta}_j^T$ is the estimate from (3.1) and $\underline{d} < 0$ and $\bar{d} > 0$ are the lower and upper excluded distance bins below and above the distance bin $j = 0$. The counterfactual distribution is then obtained by plotting $\sum_{k=0}^6 \hat{\beta}_k(j)^k$ against the distance bins.

Panel B of Figure 3.4 shows the bunching results. I estimate an average re-timing response of 18 days across treated counties and a re-timing horizon of 57 days in response to the 0.72pp tax change.⁹ Estimated bunching mass below the notch is $\hat{b} = 262$ transactions in the average treated county. I further estimate the corresponding percentage change in the average daily number of transactions—due to re-timing only—in response to the tax change, estimated for pre- and post-notch regions. Conceptually, the counterfactual world in this setting is the same time period around the reform absent re-timing responses (but not absent extensive responses). Estimation is carried out by running

$$\begin{aligned} \ln n_{it} = & \left(\sum_{k=1}^6 \beta_k^T(j)^k + \gamma_{\text{pre}}^T \mathbf{I}_{\underline{d} \leq d_{it} < 0} + \gamma_{\text{post}}^T \mathbf{I}_{0 \leq d_{it} < \bar{d}} \right) \times \text{Treat}_{it} \\ & + \sum_{k=1}^6 \beta_k(j)^k + \gamma_{\text{pre}} \mathbf{I}_{\underline{d} \leq d_{it} < 0} + \gamma_{\text{post}} \mathbf{I}_{0 \leq d_{it} < \bar{d}} + \eta \text{Treat}_{it} + \sum_{k=1}^7 \gamma_k \mathbf{I}_{\text{dow}_{it}=k} \\ & + \sum_{l=1}^{31} \delta_l \mathbf{I}_{\text{dom}_{it}=l} + \sum_{m=1}^{12} \theta_m \mathbf{I}_{\text{month}_{it}=m} + \mu \text{Holiday}_{it} + \alpha_i + \varepsilon_{it} \end{aligned} \quad (3.3)$$

where γ_{pre}^T and γ_{post}^T estimate how much the market was stimulated during the 9 days leading to the reform and subsequently depressed during the 57 days after. I find that re-timing responses increased market activity by 76% in the days prior to the reform and decreased it by 31% in the longer missing period relative to a counterfactual with extensive responses only, these effects being significant at 1%.

3.3.2 Bunching at the Border Notch

I study whether static geographic tax wedges induced by the phasing-in of the reform led to significant bunching at treatment borders. To do so, I focus on year 2015 during which groups

⁹Average re-timing is obtained by computing the average distance to the reform between $d = 0$ and $d = \bar{d}$ weighted by the missing mass at each d . Re-timing horizon corresponds to $d = \bar{d}$.

1 to 5 are always treated while group 6 and 7 are always untreated. I exclude months affected by re-timing responses, i.e., January, February and December. Treatment status differences in 2015 potentially prompted some buyers to locate just beyond treatment borders in groups 6 and 7 to avoid the extra tax liability in groups 1 to 5.

I define for each municipality its distance to the nearest treatment border, simply expressed in municipality count. A treated municipality right at a 2015 treatment border — i.e., that has at least one untreated direct neighbor — is at distance 0, all of its treated direct neighbors not also at the treatment border are at distance 1, etc. Symmetrically, an untreated municipality right at the treatment border is at distance -1 , its direct untreated neighbors not also at distance -1 are at distance -2 , etc. These distances also apply in placebo years 2010–2013.

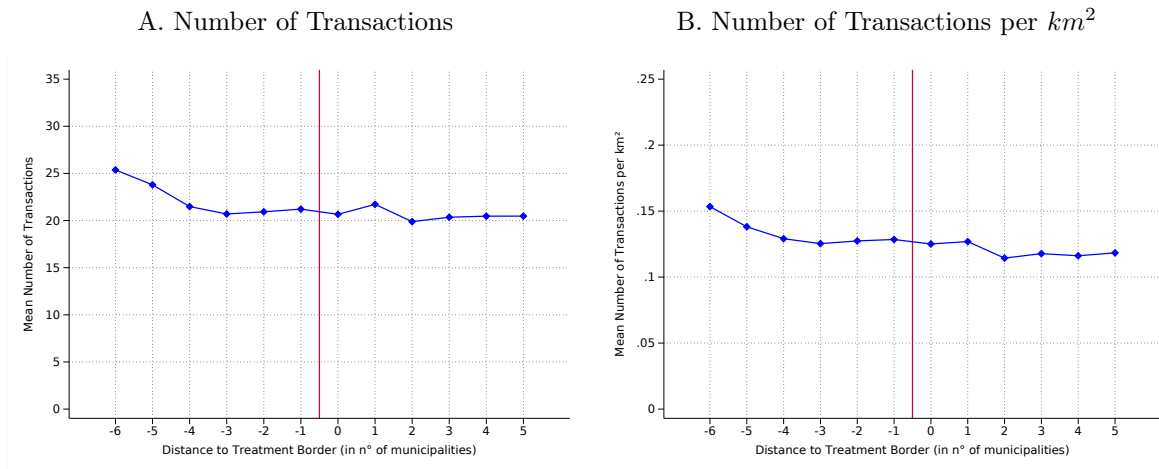
I run the following DiB specification:

$$n_{it} = \sum_{k=-7}^6 \beta_k \lambda_k + \eta \mathbb{1}_{t=2015} + \sum_{k=-7}^7 \beta_k^T \lambda_k \times \mathbb{1}_{t=2015} + \alpha + \varepsilon_{it} \quad (3.4)$$

where n_{it} is municipality i 's number of transactions in year $t \in \{2010, 2011, 2012, 2013, 2015\}$ and β_k^T is the effect of being k municipalities away from a treatment border in 2015. Distances $k = -7$ and $k = 6$ contain all municipalities below and above -7 and 6 included.

Figure 3.5 reports estimates for $\hat{\alpha} + \hat{\beta}_k^T$. As in standard I do not report estimates for $k = -7$ and $k = 6$. Panel A shows results for the number of transactions while Panel B shows

Figure 3.5: Bunching at the Border Notch



Note: This Figure shows DiB estimates for bunching at the treatment border. The dependant variable in Panel A is number of transaction. The dependant variable in Panel B is the number of transactions per km^2 .

results for the number of transactions per km^2 . Overall, Figure 3.5 points to the absence of significant bunching at treatment borders, suggesting that the tax increase was too modest to overcome buyers' location preferences.

3.4 Extensive Responses

Section 3.3 documented that the reform was salient and led to significant re-timing responses. However sharp, these responses are rather limited in time and unlikely to represent significant welfare losses. The deadweight loss is better captured by extensive responses, i.e., by the amount of mutually beneficial transactions that were kept at bay because of the tax increase. To estimate extensive responses and associated price adjustments, I turn to a DiD analysis and study how treated jurisdictions were affected by the tax change compared to untreated ones.

I exploit the phasing-in of the 2014 reform shown in Figure 3.2 to construct treatment and control units. As can be seen in Figure 3.1, treatment groups are large geographic units. In order to implement a regional policy type analysis harnessing richer spatial heterogeneity, I break each treatment group down into smaller geographic units. Several readily available geographic partitions can be chosen from. I choose municipal federations for which I report administrative boundaries in Figure 3.1. This spatial decomposition finds a middle ground between the small municipalities and the large counties. Municipal federations offer several advantages such as being spatially nested in treatment groups as well as being geographically, economically and politically relevant.¹⁰ For the rest of the paper, I aggregate all variables at the MF level.

3.4.1 Treatment, Control and Spillovers

I define treated units as the set of MFs belonging to county groups 1, 2, 3 and 4 respectively implementing the tax increase on March 1st, April 1st, May 1st and June 1st 2014. As shown in Figure 3.1, groups 1 and 2 are by far the two largest in terms of member MFs. In group 1, I drop the entire three counties neighboring Paris.¹¹ Since the first Grand Paris Express official announcement in 2009, the area has undergone massive investments in urban rail transit which

¹⁰MF membership changes over time towards more integration. I only need MFs for the geographic partition of the territory they provide. I use MFs in their 2014 form and assign membership to municipalities so that my geographic decomposition of mainland France stays constant over the period 2010–2016.

¹¹Namely *Hauts-de-Seine*, *Seine Saint-Denis* and *Val-de-Marne*.

makes it a clear outlier regarding economic and housing market dynamics.

Control units are defined as the set of MFs that experience a tax increase no sooner than January 2016, i.e., belonging to groups 6 or 7. In group 6, I drop the Paris municipality (a MF in itself) for the reasons given above. After treatment, I thus compare groups 1, 2, 3 and 4 to groups 6 and 7 according to Figure 3.2. I dismiss group 5 MFs either as control or treated units because their treatment date is too close to that of groups 1–4 or group 6.

The objective of this section is to compare market activity between different exchange economies under different tax regimes, as if they were closed jurisdictions not interacting with each other. As evidenced in Section 3.3, increasing stamp duties also prompts re-timing responses, i.e., spillovers from treated periods to untreated periods. This anticipation bias is easily dealt with by re-defining treatment dates to neutralize zero-sum re-timing responses. The possibility of transaction re-location across borders may also introduce spatial spillover biases when estimating extensive responses. However and as evidenced in Section 3.3, the reform did not seem to prompt spatial optimization responses.¹²

3.4.2 A Synthetic Control Approach

The magnitude of the reform being small—overall transaction tax increased 0.72pp from 5.09% to 5.81%—the key empirical challenge for estimating extensive responses is to find counterfactuals that very closely resemble treated units. Indeed, even modest selection into treatment is likely to be of the same order of magnitude as the treatment effect. First developed by Abadie and Gardeazabal (2003) and later enriched in Abadie et al. (2010) and Abadie et al. (2015), the synthetic control approach is well suited for studying regional effects of policy changes. It accommodates arbitrary geographic partitions, which is useful if one wants to exploit locational heterogeneity to refine the construction of counterfactuals. In my setting, it is also a convenient way to non-parametrically deal with an outcome—the number of transactions—which is bounded from below by zero. I use the synthetic control method to build an appropriate counterfactual for each MF in groups 1 to 4, using MFs in groups 6 and 7 as the “donor pool” of control units.

¹²As a robustness check, I drop border municipalities—defined as municipalities that are adjacent to at least one municipality in another treatment group—before aggregating data at the MF level. The following results are unchanged.

Matching and over-identification characteristics Let $X = \{Y, Z\}$ be a vector of characteristics attached to each MF (treated or control). Vector Y is MF pre-reform path of fortnightly number of transactions.¹³ As mentioned in Section 3.2, two nation-wide reforms uniformly increasing taxes on property capital gains were implemented on January 1st 2011 and February 1st 2012. Figure 3.3 provides evidence that they strongly affected housing markets by creating large re-timing incentives. Although synthetic counterfactuals should replicate treated units' transactions trajectory, including responses to nation-wide capital gains tax changes, it may be too ambitious an expectation in this context. In addition, Section 3.3 shows strong bunching responses prior to reform dates which are expected to bias the search for counterfactuals. While March 1st, April 1st, May 1st and June 1st 2014 are the actual reform dates for groups 1–4, it is useful to introduce the “intent-to-treat” (ITT) reform dates January 1st, February 1st, March 1st and April 1st 2014 which allow for a buffer of two months before the actual treatment dates. All together, Y contains the transaction path starting April 1st 2012—to avoid matching on responses to changes in capital gains taxation—and ending just before ITT treatment dates, i.e., ending on December 16th 2013, January 16th, February 16th or March 16th 2014—to avoid matching on re-timing responses.

Vector Z contains 2012 MF population, number of housing units and total m^2 housing surface, 2012 and 2013 mean annual household income and 2012 share of homes that are owner-occupied. It also contains 2012 and 2013 yearly mean m^2 transaction price and mean exchanged housing m^2 surface. Finally I also include 2008 educational composition on a 1–4 scale. I do not use Z in the synthetic control method but keep it for sample selection below and later over-identification checks.

Pre-selection of treated units (common support) Let $C(i)$ be the set of control MFs that can be used to build a counterfactual for MF i . Here $C(i) = C$ and is comprised of all MFs present in control groups 6 and 7. I follow suggestions in Abadie et al. (2010) and Gobillon and Magnac (2016) and find a synthetic control only to treated MFs that are not too far from control units. Specifically I only keep treated units for which all characteristics that I observe belong to the support of control units, i.e., only the i 's for which

¹³The first fortnight of each month contains its first 15 days, and the second fortnight the remaining days so that each fortnight either starts on the 1st or 16th of each month and no fortnight overlaps a reform date. If the MF belongs to control group 6 or 7, its reform date is that of group 1, 2, 3 or 4 depending on which group the treated MF to which I am finding a synthetic counterfactual belongs to.

$X_i(k) \in [\min_{j \in C(i)} X_j(k), \max_{j \in C(i)} X_j(k)]$ for each element k of X . This conservative restriction mitigates biases that arise with the synthetic control method when treated and control units have supports that are far apart.

Synthetic counterfactuals For each pre-selected treated MF i , the synthetic control procedure chooses the set of weights $\{\omega_j(i)\}_j \in [0, 1]^{\text{card}\{C(i)\}}$ in order to

$$\min_{\{\omega_j(i)\}_j} \left(Y_i - \sum_{j \in C(i)} \omega_j(i) Y_j \right)' W_i \left(Y_i - \sum_{j \in C(i)} \omega_j(i) Y_j \right) \quad (3.5)$$

where Y contains the path of pre-reform fortnightly number of transactions. This makes 42 matching periods for units of group 1 (from April 1st 2012 to December 16th 2013 included), 44 periods for units of group 2 (up to January 16th 2014), 46 for group 3 (up to February 16th 2014) and 48 for group 4 (up to March 16th 2014). The weight matrix W_i is chosen to minimize the pre-reform mean square error between i and its synthetic counterfactual outcome—here the number of transactions—as in [Abadie and Gardeazabal \(2003\)](#). I check that the number of matching periods is smaller than the number of control units $\text{card}\{C\} = 61$ to avoid small sample problems.

For each treated MF i , the procedure yields weights to apply to candidate control MFs in $C(i)$. One obtains a synthetic counterfactual, for which outcome at each distance to the reform t is

$$\hat{y}_{it} = \sum_{j \in C(i)} \omega_j(i) y_{jt} \quad (3.6)$$

where y can be the number of transactions or other MF fixed characteristics (in which case subscript t can be omitted).

Post-selection of treated units It may be useful to reward treatment-synthetic pairs for which the synthetic control best replicates the treated unit. I define the distance between i and its synthetic counterfactual as the pre-reform root mean square error:

$$d_i = \left[\frac{1}{T} \sum_{t < 0} \left(\frac{n_{it} - \hat{n}_{it}}{\bar{n}_i} \right)^2 \right]^{1/2} \quad (3.7)$$

where n_{it} is the number of transactions in the treated MF at t , \bar{n}_i is its average over the T pre-reform periods and \hat{n}_{it} is the number of transactions in the synthetic counterfactual MF

Table 3.1: Balancing Over-Identification Checks

	Treated	Synthetic	Treated	Synthetic
	Unweighted		Weighted	
Share Services	0.227	0.207	0.230	0.212
Share < High School	0.077	0.078	0.077	0.075
Share High School Pro.	0.120	0.123	0.117	0.121
Share High School Gen.	0.067	0.068	0.067	0.068
Share College	0.094	0.087	0.097	0.092
Share Owner Occ.	0.697	0.735	0.675	0.719
Population (2012)	29,466	28,863	43,538	42,994
Housing Units (2012)	12,477	12,417	18,562	18,600
Total Housing m^2 (2012)	1,127,233	1,110,758	1,642,283	1,633,865
Household Income (2013)	24,263	23,743	24,514	23,890
Average m^2 Price (2013)	1,524	1,384	1,619	1,476
Average m^2 Surface (2013)	90.53	87.39	89.10	86.32
Household Income (2012)	23,603	22,978	23,869	23,149
Average m^2 Price (2012)	1,554	1,379	1,648	1,480
Average m^2 Surface (2012)	89.70	87.47	88.21	85.85
Observations	856	856	834	834

Note: This Table reports descriptive statistics for the average treated unit and its average synthetic counterfactual. Column 1 and 2 are for simple averages while columns 3 and 4 are for match-quality weighted averages. Groups 1–4 vs. groups 6–7.

at t . Before averaging across treatment-synthetic pairs, each pair—indexed by its treated unit’s index i with a slight abuse of notation—is attributed a match quality weight which is proportional to the inverse distance between the treated MF and its synthetic counterfactual:

$$\Omega_i = \frac{1}{d_i} \left(\sum_{i'} \frac{1}{d_{i'}} \right)^{-1} \quad (3.8)$$

so that $\sum_i \Omega_i = 1$. For any outcome y —transaction trajectory or MF characteristics—I compute its weighted average across the many pairs of treated and synthetic MFs:¹⁴

$$\begin{aligned} y_t &= \sum_i \Omega_i y_{it} \\ \hat{y}_t &= \sum_i \Omega_i \hat{y}_{it} \end{aligned} \quad (3.9)$$

¹⁴When y is the number of transactions, because I exploit different reforms happening at different dates, I center each treated MF and its synthetic counterfactual around its reform date. Average treatment and synthetic outcome series are reported in terms of distance t to the reform: I trim top and/or bottom distances when necessary to make sure that the set of treatment-synthetic pairs contributing to the averages is the same at each t .

Treatment vs. synthetic Table 3.1 reports summary statistics for the vector of characteristics Z —which was not used to build synthetic counterfactuals—for mean treated unit and mean synthetic counterfactual. It provides over-identification checks for unweighted and weighted averages. The two columns show that the synthetic control method performs well in replicating a battery of pre-reform characteristics. Importantly, this intermediate result is obtained as a by-product of matching only on the pre-trends.

3.4.3 Average Effect on Number of Transactions

I express average treatment and synthetic control transaction trajectories in both absolute and relative terms. For the latter I normalize the two series n_t and \hat{n}_t by their pre-reform average value. The final treatment effect obtained when comparing my normalized average treated MF and its normalized average synthetic counterfactual is thus the percentage change in the number of transactions due to the reform.

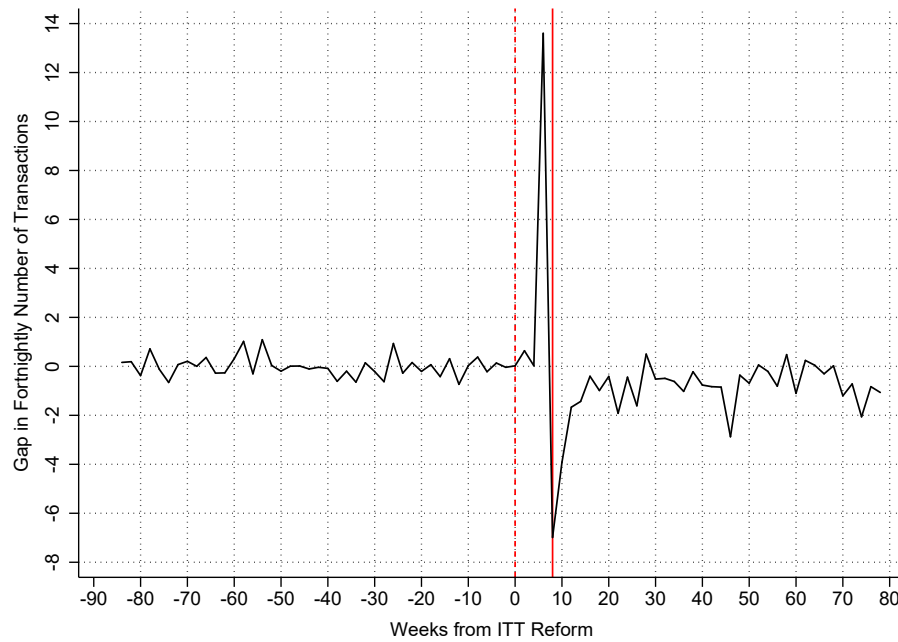
Figure 3.6 reports the transaction gap between average treated and average synthetic control. Panel A of Figure 3.7 reports the gap in normalized transaction path. Panel B only keeps treatment groups 1 and 2 to retain more temporal depth after treatment. Panels C and D are similar to Panels A and B, except that the synthetic control procedure is carried out using only group 7 as the control group to get as many post-treatment periods as possible. Taken together, these Figures suggest a modest long-lasting decrease in transaction volume after the 0.72pp tax increase somewhere between 0 and 10% of the counterfactual volume. Furthermore, the underlying post-reform gap looks stable in Panels C and D that show more post-treatment periods. This is a first piece of evidence that the apparent effect of the reform is not driven by chance, that is, not coming from the synthetic control’s poor out-of-sample performance. However, the noise-to-signal ratio is quite high and one needs further manipulations to precisely quantify extensive responses.

In order to smooth errors out, I turn to the cumulative number of transactions. The synthetic control procedure is unchanged up to step (3.9) included, that I now supplement with

$$\begin{aligned} N_t &= \sum_{d \leq t} n_d \\ \hat{N}_t &= \sum_{d \leq t} \hat{n}_d \end{aligned} \tag{3.10}$$

which are treated and synthetic cumulative number of transactions up to distance t to the

Figure 3.6: Gap in Number of Transactions



Note: This Figure shows the synthetic control estimation result for the raw number of transactions. The solid red line shows the actual reform time, while the dotted red line shows an “intent-to-treat” (ITT) reform time that accounts for re-timing responses when finding synthetic counterfactuals to treated units. Treated MFs belong to groups 1–4 while control “donor” MFs belong to groups 6–7. Average treated and synthetic number of transactions are computed at each distance to the ITT reform after centering each treated unit and its synthetic counterfactual around its reform date, and dropping top and bottom distances when necessary to force a balance panel in terms of distance to ITT reform time.

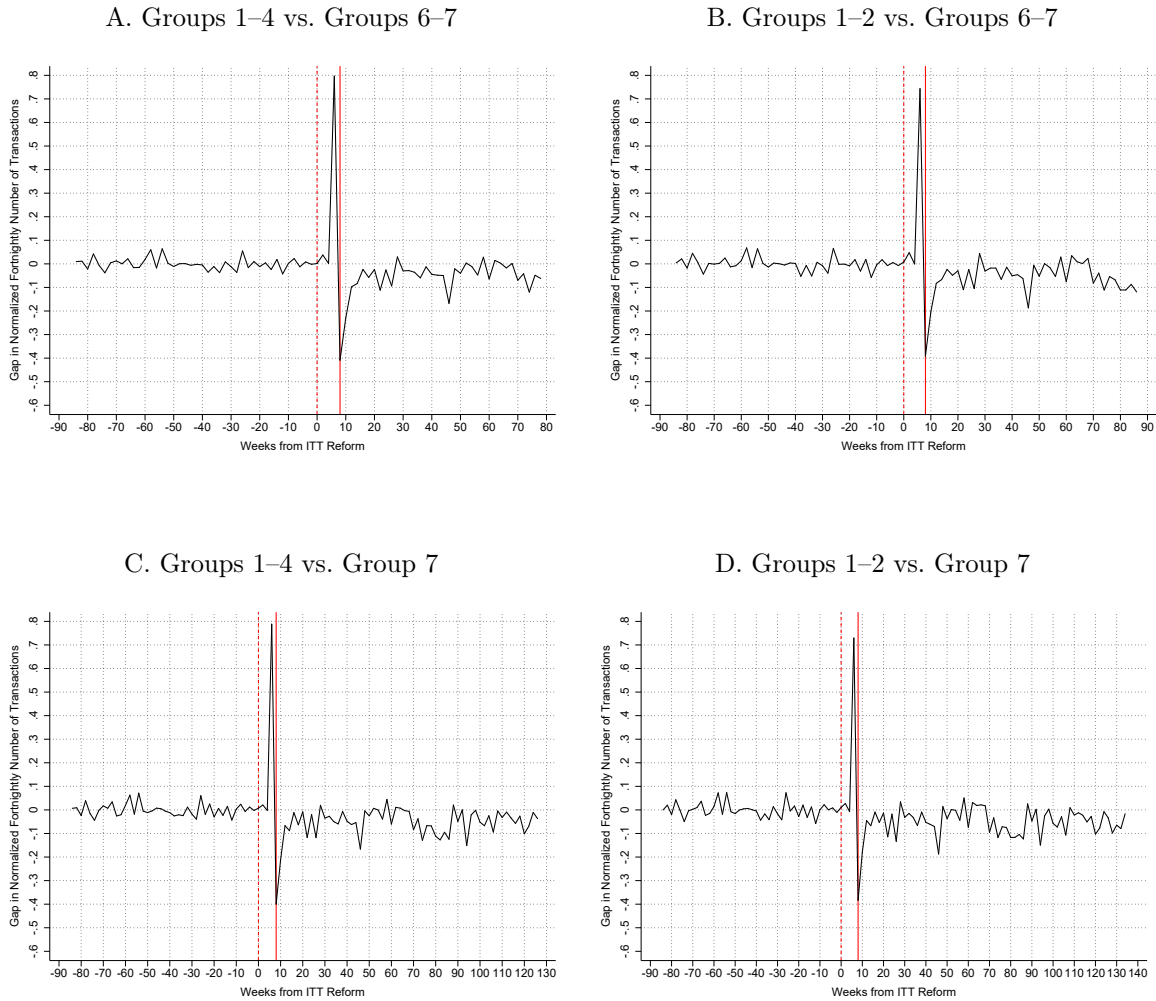
reform.¹⁵ If the synthetic control method is successful in replicating treated units’ absolute transaction count, it should also be in replicating treated units’ cumulative transaction count, starting from any arbitrary date. The cumulative difference $N_t - \hat{N}_t$ should smooth out fitting errors between the average treated unit and its average synthetic counterfactual if they do not capture a systematic bias. Importantly, $N_t - \hat{N}_t$ should also gradually absorb re-timing responses, since their pre- and post-reform contributions to the number of transactions sum to zero.¹⁶

Figure 3.8 reports the gap in cumulative number of transactions $N_t - \hat{N}_t$. Panel A is

¹⁵I start from the maximum distance to the reform which is common to all treated groups so that my panel of treatment-synthetic pairs is balanced in terms of distance t to the reform.

¹⁶Taking the cumulative sum of log transactions, although useful to directly interpret effects in relative terms, is ill-advised here because of re-timing responses. Indeed, re-timing responses and the concavity of the log ensure that the final cumulative difference between a treated unit and a perfect counterfactual will be negative even in the absence of extensive responses.

Figure 3.7: Gap in Normalized Number of Transactions

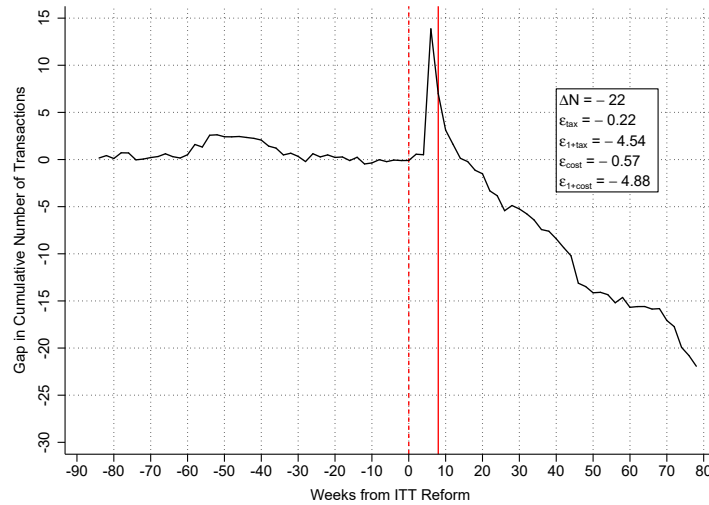


Note: This Figure shows synthetic control estimation results for normalized number of transactions. Panel A shows groups 1–4 against groups 6–7. Panel B shows groups 1–2 against groups 6–7. Panels C and D offer more post-treatment temporal depth by showing groups 1–4 and groups 1–2 against group 7. Treated and control units as well as synthetic matching weights and match quality weights are taken from the estimation for raw number of transactions. The only change is the ex-post normalization of average treated and average synthetic trajectories by their pre-ITT-reform average.

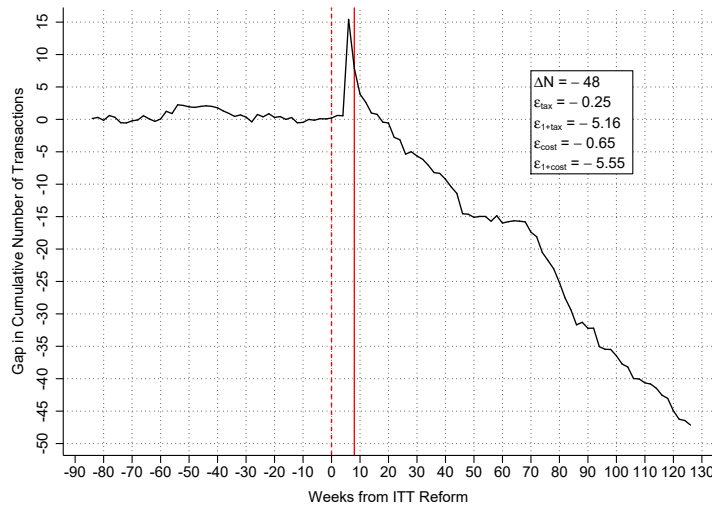
the main specification using treatment groups 1–4 and control groups 6–7 while Panel B shows 1–4 versus 7 only. The cumulative sums eliminate a lot of noise in the data which confirms that my synthetic counterfactuals adequately replicate treated units’ fundamental transaction path. Figure 3.8 offers compelling evidence that the reform had long-lasting effect on extensive margin responses since the cumulative gap steadily increases. The last point $N_T - \hat{N}_T$ in Panel A of Figure 3.8 is the causal effect of a 0.72pp tax increase on the total

Figure 3.8: Gap in Cumulative Number of Transactions

A. Groups 1–4 vs. Groups 6–7



B. Groups 1–4 vs. Group 7



Note: This Figure shows synthetic control estimation results for the cumulative number of transactions. In Panel A, treated MFs belong to groups 1–4 while control “donor” MFs belong to groups 6–7. Panel B shows similar estimates for groups 1–4 against group 7. Treated and control units as well as synthetic matching weights and match quality weights are taken from the estimation for raw number of transactions. The only change is that I take the cumulative sum of transactions for average treated unit and its average synthetic counterfactual starting from the minimum available distance to ITT reform common to all treated groups.

number of transactions between $t = 0$ and $t = T$ due to extensive responses. I estimate a loss of $N_T - \hat{N}_T = -22$ transactions in the average treated municipal federation over 20 months

with the main specification.

As evidenced in Figure 3.8, absolute treatment effects depend on the estimation specification. Indeed, sample pre- and post-selection criteria influence average treated and synthetic baseline market sizes. To express treatment effect in percentage terms, I compute the *synthetic control's* total number of transactions between $t = 0$ and $t = T$, that is, $\hat{N}_T - \hat{N}_{-1}$. Indeed, I cannot compare $N_T - \hat{N}_T$ to total post-reform number of transactions in the average treated unit since the latter is contaminated by re-timing responses. However, I use the fact that $N_T - \hat{N}_T$ would have been the amount of extensive responses in the average synthetic control *if it had been treated* since synthetic counterfactuals achieves replication of treated units in levels. I obtain a relative treatment effect of $\frac{\Delta n}{n} = -3.1\%$, i.e., the tax increase depressed market activity by an average 3.1% in treated areas, assuming that this relative treatment effect can be considered constant over time and across locations. This percentage change lies in the range of existing estimates studying the effect of the same reform: Bachelet and Poulhès (2017) find $\frac{\Delta n}{n} = -10\%$ while Bérard and Trannoy (2018) find $\frac{\Delta n}{n} = 0\%$ for medium-term extensive responses.

Elasticities and comparison with existing results I use this estimated relative effect to compute the elasticity of transactions to the net tax rate $\varepsilon_{n,\tau} \approx -\frac{\Delta n}{n} \frac{\tau}{\Delta\tau} = 0.22$ which I will use in the welfare application. This elasticity captures modest reduced-form behavioral responses to transaction taxes. To compare my estimates to existing results in other settings, I compute some additional statistics. The elasticity of transactions to the gross tax rate is $\varepsilon_{n,1+\tau} \approx -\frac{\Delta n}{n} \frac{1+\tau}{\Delta\tau} = 4.54$. This second elasticity captures behavioral responses to publicly imposed gross price increase. As a comparison point, Best and Kleven (2017) find $\varepsilon_{1+\tau} = 14.3$ for the UK, suggesting that the French housing market may be stickier than the UK market. Assuming extra moving-related costs of $c_m = 8\%$, one can also compute the elasticity to net total moving costs $\varepsilon_{\text{cost}} \approx -\frac{\Delta n}{n} \frac{c_m+\tau}{\Delta\tau} = 0.57$.¹⁷ This estimate is smaller than the elasticity of $\varepsilon_{\text{cost}} \approx 1$ that Dachis et al. (2011) find for Toronto.

Taken together, these elasticities suggest that French households are somewhat sensitive to changes in transaction costs. In Panel B of Figure 3.8, I estimate a different absolute effect using only group 7 as the control group. However, elasticities are largely unchanged compared to the main specification of Panel A.

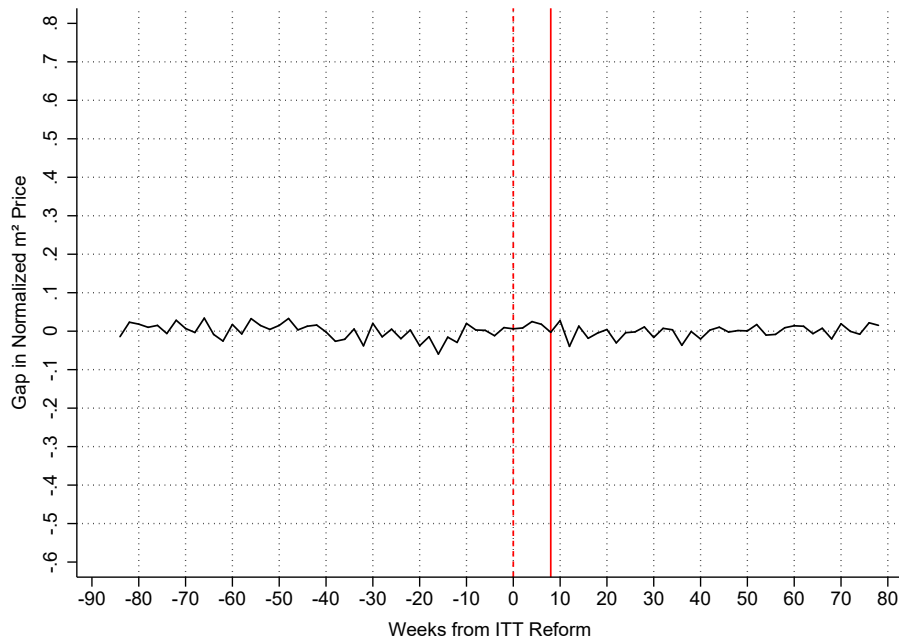
¹⁷Notary fees are typically around 2%, and we assume that other moving costs represent an extra 6% (e.g., agency fees, professional movers, etc.) as in Dachis et al. (2011).

3.4.4 Average Effect on Prices and Quality

I investigate how the reform affected prices and the quality composition of exchanged homes. Importantly, the synthetic control procedure is unchanged up to equation (3.9) included. That is, I match on transaction pre-trends and weight each treated-synthetic pair by transaction goodness of fit defined in equation (3.8). Thus, the estimation sample for prices or quality is the same as that used for estimating extensive responses above. The only change is that average treated and synthetic outcomes are respectively normalized with respect to mean pre-reform price or quality.

Figure 3.9 reports the results for m^2 prices. Strikingly, the reform had no effect. A standard interpretation for this result is that the supply of existing homes is very elastic, even in the short run, so that tax incidence fell on buyers. This interpretation may however be partly biased by quality or composition effects. Think of a model with discrete and ordered quality for homes and a continuum of households ordered by income. In such a model, there

Figure 3.9: Gap in Normalized m^2 Price

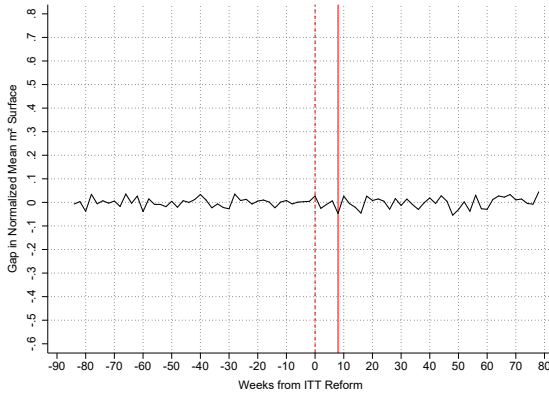


Note: This Figure shows the synthetic control estimation result for normalized m^2 prices. Treated MFs belong to groups 1–4 while control or “donor” MF belong to groups 6–7. Treated and control units as well as synthetic matching weights and match quality weights are taken from the estimation for raw number of transactions.

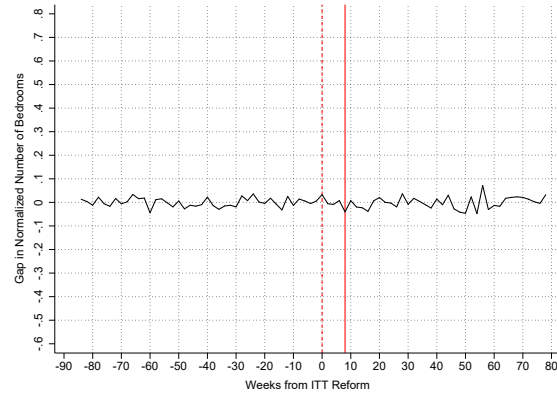
will be perfect sorting, i.e., the tail of the income distribution will buy top quality homes, the second richest group homes of second-best quality, etc. The poorest group cannot afford to buy, while the second poorest buys lowest-quality homes. In equilibrium, each “quality” market will host marginal buyers that are indifferent between this home quality and that just below. A small tax increase will induce a chain of substitution towards lower quality homes from these marginal buyers, driving down average quality. However buyers of lowest quality homes who end up not buying following the tax increase will drive up average exchanged

Figure 3.10: Quality Changes

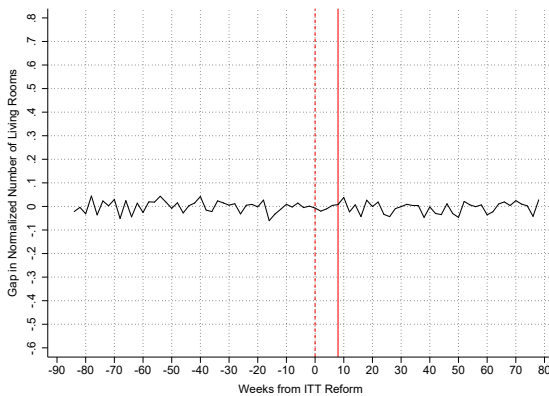
A. Gap in Normalized Mean m^2 Floor Surface



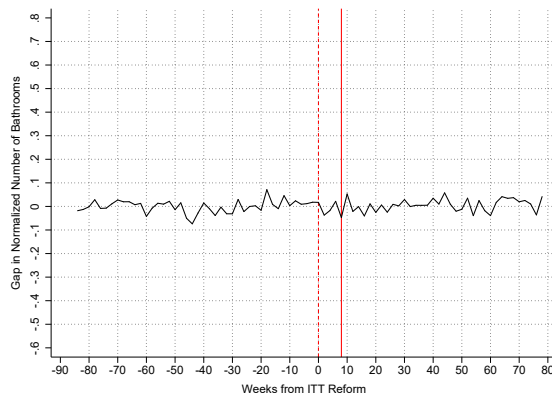
B. Gap in Normalized Number of Bedrooms



C. Gap in Normalized Number of Living Rooms



D. Gap in Normalized Number of Bathrooms



Note: This Figure shows synthetic control estimation results for normalized home quality measures. Treated MFs belong to groups 1–4 while control “donor” MFs belong to groups 6–7. Treated and control units as well as synthetic matching weights and match quality weights are taken from the estimation for raw number of transactions.

quality. Assuming away general equilibrium price adjustments, there will hence be a direct or micro effect of the tax increase on average purchase price that will only transit through this quality channel. If density is high enough at the lowest income cutoff where buyers are indifferent between buying lowest-quality homes and not buying—so that average exchanged quality actually goes up—the overall price change may even be positive.

Figure 3.10 explores quality changes, and dismisses the hypothesis that composition plays an important role here. Indeed, mean home surface, number of living rooms, bedrooms and bathrooms are unaffected by the tax increase, suggesting the need for another mechanism to explain why prices are apparently unchanged. As mentioned, a very elastic housing supply may be behind price responses. In Section 3.5, I explore an alternative mechanism based on buyers' heterogeneity in a search framework.

3.4.5 Statistical Inference

Statistical inference investigates whether the results of the above sections are driven by chance. The synthetic control method could yield faithful replicas of treated units in-sample—that is, for pre-reform periods—while not guaranteeing out-of-sample replication. As shown by [Abadie et al. \(2015\)](#), statistical inference would become unnecessary if I was able to match trajectories on an infinite number of fixed-length pre-reform periods, since I would obtain a perfect counterfactual.

One could typically employ cross-validation or other model selection methods (lasso, etc.) to help improve out-of-sample predictive power. Here, one would choose the number of matching periods as well as the importance weight matrix W_i to strike a balance between systematic bias—how far off the synthetic counterfactual is from the treated unit on average—and the variance of the estimated effect. Such methods would serve to improve confidence in my synthetic counterfactuals.

As a first and maybe sufficient check, I can measure the extent to which the method outlined in Section 3.4 tends to yield proper counterfactuals with adequate out-of-sample behaviors. These are the so-called placebo tests. The following method is adapted from [Cavallo et al. \(2013\)](#). For each MF i :

1. Fit a synthetic control method on each control unit in the donor pool $C(i)$ according to equation (3.5), as if that control unit were treated on the same day as i while the remaining units in $C(i)$ are considered the new donor pool.

2. Store the outcome path of each control unit of $C(i)$ as well as that of its synthetic counterfactuals in a matrix M_i .¹⁸

Then, do the following 10,000 times:

3. Randomly draw one control-synthetic pair from each M_i . Compute the average control and the average synthetic trajectory, potentially weighting each pair by a match quality weight as in 3.4.2.
4. Compute the cumulative sum of the average control's and average synthetic's transaction path. Estimate a placebo treatment effect in percentage using the method in 3.4.3. This average normalized placebo effect can be compared to my average normalized treatment effect. With this method, each placebo effect is computed using the same number of observations as when estimating the treatment effect.

Table 3.2: Balancing Over-Identification Checks (Average Placebo)

	Treated	Synthetic	Treated	Synthetic
	Unweighted		Weighted	
Share Services	0.217	0.210	0.221	0.215
Share < High School	0.075	0.076	0.073	0.074
Share High School Pro.	0.123	0.122	0.121	0.119
Share High School Gen.	0.070	0.068	0.069	0.068
Share College	0.095	0.090	0.099	0.095
Share Owner Occ.	0.711	0.727	0.697	0.712
Population (2012)	34,029	31,698	47,349	44,630
Housing Units (2012)	14,511	13,532	20,370	19,139
Total Housing m^2 (2012)	1,352,063	1,220,127	1,887,454	1,704,731
Household Income (2013)	24,234	23,966	24,649	24,308
Average m^2 Price (2013)	1,460	1,421	1,560	1,515
Average m^2 Surface (2013)	89.22	87.17	88.98	86.60
Household Income (2012)	23,497	23,220	23,918	23,576
Average m^2 Price (2012)	1,471	1,434	1,575	1,533
Average m^2 Surface (2012)	87.70	86.78	87.79	85.95
Observations	172	172	172	172

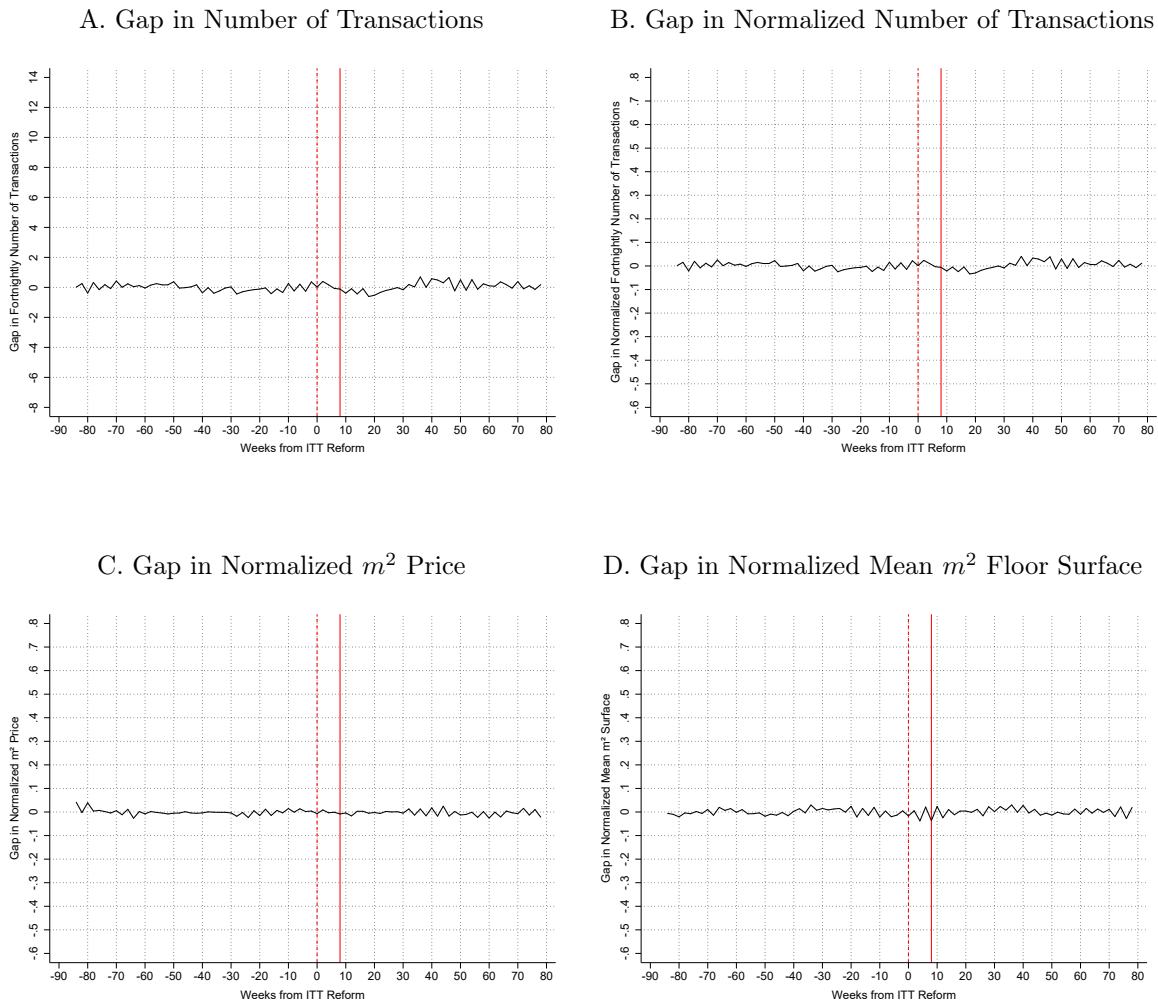
Note: This Table reports descriptive statistics for the average control “donor” unit and its average synthetic counterfactual. Column 1 and 2 are for simple averages while columns 3 and 4 are for match-quality weighted averages. Control units belong to groups 6–7.

¹⁸If all treated units are treated at the same time, M_i will be the same for all treated units. When the donor pool is the same for all treated units, this method amounts to drawing placebo effects with replacement.

With step 4 carried-out 10,000 times, I compute the mean of the distribution of average control-synthetic gap for each outcome. Figure 3.11 reports these mean gaps. The synthetic control method performs well after placebo reforms for replicating the path of control units. Table 3.2 further shows that mean control and synthetic units are very close in terms of observable characteristics.

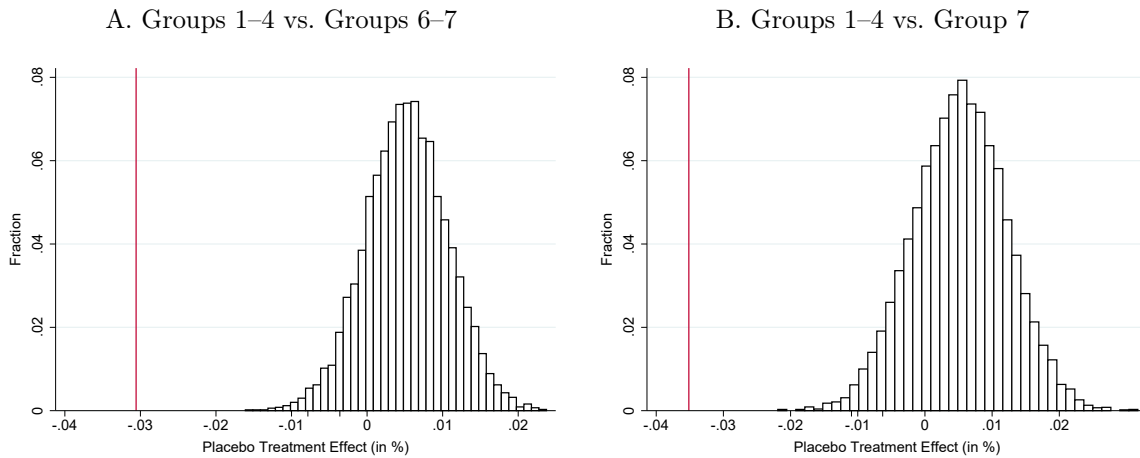
With these 10,000 iterations and step 4, I get a distribution of average normalized placebo effects for extensive responses and can extract percentile-based rejection cutoffs. I compare my treatment effect to these rejection cutoffs. Figure 3.12 reports the distribution of placebo

Figure 3.11: Average Placebo Estimates



Note: This Figure shows synthetic control estimation results for average placebo effects. All MFs belong to groups 6–7.

Figure 3.12: Treatment vs. Placebo Effects



Note: This Figure shows the distribution of placebo effects. Each placebo effect is obtained by running the full synthetic control procedure described in 3.4.2 on a bootstrap sample of control “donor” units. The distribution is obtained from 10,000 bootstrap replications. The red vertical line shows the treatment effect estimate. Groups 1-4 against groups 6-7 in Panel A and against group 7 in Panel B.

percentage drops in the fortnightly number of transactions as well as the 1% rejection cutoff, along with the original estimate of Figure 3.8. First, the mean of the distribution is close to zero showing that the method performs well also for the cumulative number of transactions. Second, my estimated treatment effect of -3.1% is highly significant and actually higher in absolute value than all placebo effects.

3.4.6 Public Spending

Tax proceeds are not thrown away and one might argue that my estimates capture the effect of increased public good provision in the counties raising their taxes. From this angle, the advantage of the 2014 reform is that it applied uniformly within large geographic areas that may be considered closed markets from the perspective of residential choices. If public goods enter utility separably from traded goods and housing consumption, additional tax proceeds can be thought of as lump-sum increases in local public services. Hence, transaction taxes tax can be considered as purely distorting the allocation of property between residents of a same county group. In practice, French counties are mainly in charge of means-tested redistribution which is expected to increase housing consumption through income effects, so that my “policy elasticity” estimate may actually be a lower bound for behavioral responses. Finally, stamp duty increases were introduced to compensate budgetary cuts to local governments, which

suggests that local budgets were almost unchanged by the reform.

3.5 A Simple Search Model

I illustrate the welfare consequences of a transaction tax in a simple one-off matching market. In particular, I propose a simple mechanism that rationalizes the observed price responses in Section 3.4: transaction taxes enhance match surplus by screening buyers with the lowest reservation values out of the housing market. Higher taxes have the classic depressing effect on bargained price but lead to higher matching efficiency which drives up prices, so that the overall effect is ambiguous.

3.5.1 Model

There is a continuum of measure one of risk-neutral heterogeneous buyers, indexed by i and characterized by their reservation value b_i which is smoothly distributed with cumulative density function $F(\cdot)$. There is a continuum of risk-neutral homogeneous sellers of measure one. One-sided heterogeneity greatly simplifies the setup without sacrificing the core intuition. I assume that sellers' shared reservation value is common knowledge so that buyers correctly anticipate transaction prices.

Market environment The number of buyer-seller matches is $n(\bar{e}, \bar{v}) < 1$ where \bar{e} is total search intensity among heterogeneous buyers, and \bar{v} is total sellers' search intensity. Matching technology n is differentiable, increasing in both arguments and homogeneous of degree one. I will show that all matches lead to transactions in this setup. Define market tightness as the ratio of aggregate seller effort to aggregate buyer effort $\theta = \bar{v}/\bar{e}$, where total buyer and seller search intensity result from both extensive and intensive margin choices. Buyers' and sellers' matching probability per unit of search effort are $\alpha_b(\theta) = n(\bar{e}, \bar{v})/\bar{e} = n(1, \theta)$ and $\alpha_s(\theta) = n(\bar{e}, \bar{v})/\bar{v} = n(1/\theta, 1)$ respectively, with $\alpha'_b(\theta) > 0$ and $\alpha'_s(\theta) < 0$.

Buyers Upon matching, a type- i buyer and a seller Nash-bargain over the transaction price p_i , that is

$$\max_{p_i} = \beta \ln(b_i - p_i - \tau) + (1 - \beta) \ln(p_i - s) \quad (3.11)$$

where $\beta \in [0, 1]$ is buyer's bargaining power and is independent of the buyer's type. The transaction tax is an excise tax, in contrast with real-world ad-valorem taxes. This technical

assumption eases the welfare discussion below while retaining the core intuition.¹⁹

I assume that β is exogenous so that it does not depend on relative supply and demand forces summarized by market tightness θ . While unrealistic, this assumption helps emphasize the composition channel. In this setup, housing prices only depend on tax τ and on which buyer i is matched with a seller. Transaction price for buyer i is then

$$p_i(\tau) = \beta s + (1 - \beta)(b_i - \tau) \quad (3.12)$$

The bargained price p_i increases with both buyer's and seller's reservation value b_i and s and decreases with tax τ .

As mentioned, I assume that buyers know sellers' reservation value s . Anticipating the negotiated housing price, each buyer makes the extensive choice of whether to search the housing market or not. Conditional on searching, a buyer decides how much search effort e_i to exert. Search costs $c_b(\cdot)$ are differentiable, increasing and convex with $c_b(0) = c'_b(0) = 0$. Type- i buyer's utility is

$$u_i = \begin{cases} e_i \alpha_b(\theta) \beta [b_i - \tau - s] - c_b(e_i) & \text{if searching} \\ 0 & \text{if not searching} \end{cases} \quad (3.13)$$

where I have plugged in the transaction price p_i . Conditional on searching, buyer i 's search effort $e_i(\theta, \tau)$ is implicitly defined by $c'_b(e_i) = \alpha_b(\theta) \beta [b_i - \tau - s]$. Conditional search effort increases with taste for owning property b_i , increases with match-finding probability itself increasing in tightness θ and decreases with tax τ . Buyers with $b_i = \tau + s$ search with exactly zero effort $e_i = 0$ hence are indifferent between searching and not searching. Buyers with $b_i > \tau + s$ search with strictly positive effort since the marginal benefit of searching from $e_i = 0$ is $\alpha(\theta) \beta [b_i - \tau - s] > 0$ while the marginal cost is $c'_b(0) = 0$. Buyers with $b_i < \tau + s$ do not search since searching necessarily yields $u_i < 0$.

Call $\mathcal{B}(\tau) = \{i \mid b_i \geq \tau + s\}$ the set of active buyers at policy τ with measure $1 - F(\tau + s)$. Note that $\mathcal{B}(\tau)$ does not depend on θ : tightness does not affect the number of active buyers. Total buyer search effort is

$$\bar{e}(\theta, \tau) = \int_{\mathcal{B}(\tau)} e_i(\theta, \tau) \quad (3.14)$$

Recall that the marginal active buyer exerts zero effort. Aggregate buyer search effort hence

¹⁹All arguments in this section remain valid with an ad-valorem tax.

changes with that of infra-marginal active buyers, i.e., increases with θ and decreases with τ .

Sellers Sellers behave homogeneously and have utility

$$u = \begin{cases} s + v \alpha_s(\theta) (p^e - s) - c_s(v) & \text{if searching} \\ s & \text{if not searching} \end{cases} \quad (3.15)$$

where p^e is the expected transaction price taken as given by sellers and $c_s(\cdot)$ behaves as $c_b(\cdot)$. Search will be positive only if $p^e \geq s$. Conditional search intensity $v(\theta, p^e)$ is implicitly defined by $c'_s(v) = \alpha_s(\theta) (p^e - s)$. Seller search effort decreases with θ and increases with observed mean selling price p^e . It does not directly depend on τ . Total seller search effort is

$$\bar{v}(\theta, p^e) = v(\theta, p^e) \times \mathbf{1}_{p^e \geq s} \quad (3.16)$$

Equilibrium Equilibrium is characterized by tightness and expected price $\{\theta^*(\tau), p^{e*}(\tau)\}$ satisfying

$$\begin{aligned} \theta &= \frac{\bar{v}(\theta, p^e)}{\bar{e}(\theta, \tau)} \\ p^e &= \mathbb{E}[p|\theta, \tau] = \int_{\mathcal{B}(\tau)} \frac{e_i(\theta, \tau)}{\bar{e}(\theta, \tau)} p_i(\tau) \end{aligned} \quad (3.17)$$

Note that active buyers verify $b_i \geq \tau + s$ so equilibrium price $p^e \geq s$ when $\tau \geq 0$. Hence all sellers will exert positive effort in equilibrium.

The first equation yields $\theta(p^e, \tau)$, increasing in its two arguments (see Appendix C.1). The second equation yields $p^e(\theta, \tau)$ which varies with market tightness according to:

$$\frac{\partial p^e}{\partial \theta} = \int_{\mathcal{B}(\tau)} \frac{\partial (e_i/\bar{e})}{\partial \theta} p_i \quad (3.18)$$

This relationship captures composition effects and its sign is ambiguous. It is negative if buyers with lower reservation value increase their effort proportionally more than buyers with higher reservation value when the matching probability increases. Furthermore, $p^e(\theta, \tau)$ changes with τ according to:

$$\frac{\partial p^e}{\partial \tau} = \int_{\mathcal{B}(\tau)} \frac{e_i}{\bar{e}} \frac{\partial p_i}{\partial \tau} + \int_{\mathcal{B}(\tau)} \frac{\partial (e_i/\bar{e})}{\partial \tau} p_i \quad (3.19)$$

where I can neglect changes in $\mathcal{B}(\tau)$ because the marginal buyer searches with zero effort. The first term is negative and reflects how degrading match surplus adversely affects prices. The

second term reflects composition changes and its sign is ambiguous here as well. It is positive if low types decrease their effort proportionally more than high types when taxes increase.

Assumption 1. (Low Types Are More Responsive) *The model's primitives are such that*²⁰

$$\frac{d}{db_i} \frac{\partial \ln e_i}{\partial \theta} \leq 0 \quad \text{and} \quad \frac{d}{db_i} \frac{\partial \ln e_i}{\partial \tau} \geq 0 \quad (3.20)$$

Proposition 1. (Tax May Increase Price) *When agents are homogeneous, a tax increase always decreases equilibrium price p^e . With heterogeneous agents and Assumption 1, a tax increase may increase or decrease p^e .*

Proof. See Appendix C.2. □

With heterogeneous agents and without more structure on the model, it is *a priori* impossible to know which effect will prevail. This simple mechanism may partly explain why prices do not fall as expected in Section 3.4. For completeness, I show that the model equilibrium described in (3.17) is well-defined under standard parameterizations.

Proposition 2. (Equilibrium Existence and Uniqueness) *Under standard functional forms $c(e) = e^\gamma$ with $\gamma > 1$ and $n(\bar{e}, \bar{v}) = \bar{e}^\eta \bar{v}^{1-\eta}$ with $\eta \in [0, 1]$, the equilibrium exists and is unique.*

Proof. See Appendix C.3. □

3.5.2 Welfare

Even without specifying further the model's primitives, one can still characterize the welfare effect of a tax increase with a set of sufficient statistics (Chetty 2009) that are robust to the model's unspecified deep parameters. The transaction tax is rebated lump-sum to all agents, whether participating or not in the property market. I consider the following money-metric criterion which sums agents' expected utility:

$$W = s + \bar{v} \alpha_s(\theta) (\mathbb{E}[p|\theta, \tau] - s) - c_s(\bar{v}) + \int_{\mathcal{B}(\tau)} (e_i \alpha_b(\theta) [b_i - \tau - p_i] - c_b(e_i)) + n \tau \quad (3.21)$$

Manipulation yields a simple expression featuring sellers' reservation value, gains from trade and search costs:

$$W = s + n \times (\mathbb{E}[b|\tau] - s) - c_s(\bar{v}) - \int_{\mathcal{B}(\tau)} c_b(e_i) \quad (3.22)$$

²⁰It will be the case for standard search cost functions $c_b(e) = e^\gamma$ with $\gamma > 1$, in which case $\frac{d}{db_i} \frac{\partial \ln e_i}{\partial \theta} = 0$ and $\frac{d}{db_i} \frac{\partial \ln e_i}{\partial \tau} > 0$.

Note that expression (3.22) does not feature prices as terms of trade only operate redistribution between buyers and sellers.

It is well known that in search markets with homogeneous agents, the decentralized equilibrium is generally inefficient because the price-setting mechanism does not ensure efficiency. Indeed, buyers search too much as they do not internalize that own search effort hurts other buyers' matching probability. They also search too little, not internalizing that their search effort increases sellers' probability to find a trade partner. A knife-edge situation in which these two externalities offset each other is when the so-called Hosios condition (Hosios 1990) holds, i.e., when the share that buyers (resp. sellers) extract from the match surplus equals the elasticity of the matching function with respect to their search effort.

Since I wish to emphasize the misallocation inefficiency that comes with buyer heterogeneity, I impose Hosios' condition. This allows me to focus on how search externalities affects gains from trade $\mathbb{E}[b|\tau] - s$, i.e., who gets matched with a seller.

Assumption 2. *Bargaining is efficient, i.e., the Hosios condition holds:*

$$\beta = \frac{\bar{e}}{n} \frac{\partial n}{\partial \bar{e}} = 1 - \frac{\bar{v}}{n} \frac{\partial n}{\partial \bar{v}} \quad (3.23)$$

Proposition 3. (Welfare Change) *Under Assumption 2, welfare changes with the transaction tax according to:*

$$\frac{dW}{d\tau} = \tau \frac{dn}{d\tau} + n(1 - \beta) \frac{d\mathbb{E}[b]}{d\tau} \quad (3.24)$$

Proof. See Appendix C.4. □

The proof relies on standard envelope conditions. The first term is the impact of increasing stamp duties on the number of mutually beneficial transactions. The second term captures how the tax increase affects the quality of these transactions, fully characterized here by expected buyer reservation value. Proposition 3 sheds lights on the nature of the matching inefficiency as an externality that low-type active buyers impose on sellers. Upgrading gains from trade is welfare improving only if sellers are able to capture part of the match surplus, i.e., if $\beta < 1$. When buyers extract the full match surplus—and are the only ones whose efforts matter for matching under Assumption 2—this matching externality vanishes. In a more complex model with seller heterogeneity, this welfare channel is expected to be two-sided. Sellers with higher reservation values may decrease their effort more following a tax increase, increasing buyers' probability to be matched with a low-valuation seller hence to pay a low price.

One can use Proposition 3 to give a sufficient statistics characterization of the optimal transaction tax in this stylized context:

$$\tau^* = (1 - \beta) \mathbb{E}[b] \frac{\varepsilon_{b,\tau}}{\varepsilon_{n,\tau}} \quad (3.25)$$

with $\varepsilon_{b,\tau} \equiv \frac{d\mathbb{E}[b]}{d\tau} \frac{\tau}{\mathbb{E}[b]} > 0$ under Assumption 1 and $\varepsilon_{n,\tau} \equiv -\frac{dn}{d\tau} \frac{\tau}{n}$. All variables are observed at the optimal tax so this characterization is implicit. The optimal tax trades off market activity and the (expected) gains from trade per transaction. Increasing τ decreases lower types' search effort more than higher types', improving the odds that the latter are matched hence upgrading expected match quality. However, it may cause efficiency losses by decreasing overall search effort and total number of transactions. One checks that $\tau^* = 0$ absent buyer heterogeneity ($\varepsilon_{b,\tau} = 0$).

Finally, I use Proposition 3 to carry out a back-of-the-envelope welfare assessment of the 2014 reform through the lens of the channels discussed here. Tax changes that followed the 2014 reform were small, hence providing an ideal setup for the sufficient statistics approach—the evaluation of a small past reform (see Kleven 2018b). To reduce the number of unobserved variables, one can plug prices in (3.24) according to $d\mathbb{E}[p] = (1 - \beta)(d\mathbb{E}[b] - 1)$. Hence, the welfare change from the tax increase can be re-written as

$$\frac{dW}{d\tau} = 1 - \beta - \varepsilon_{n,\tau} + \frac{d\mathbb{E}[p]}{d\tau} \quad (3.26)$$

While prices as terms of trade are absent from equation (3.22), their variation appears here as a measure of changes in allocative efficiency. In Section 3.4, I estimate $\frac{d\mathbb{E}[p]}{d\tau} = 0$ and $\varepsilon_{n,\tau} = 0.22$. Hence in this framework, increasing stamp duties improved welfare if $\beta < 0.78$, i.e., if buyers capture less than 78% of match surplus.

Complete characterization of the reform's welfare impact in this stylized setup hinges on the ability to estimate buyers' bargaining power, or equivalently the elasticity of matches to buyer search effort, an area where evidence is scant. For instance, Besley et al. (2014) calibrate a bargaining model with knowledge of sellers' reservation value and estimate $\beta \approx 0.6$ for the UK. One would expect home buyers in France, where housing supply elasticity is typically lower than in the UK, to have even less bargaining power than their UK counterparts. In the context of the 2014 stamp duty reform, the composition channel discussed above likely led to a positive effect on welfare that outweighed the deadweight loss caused by missing transactions. However, this result obtains in a deliberately illustrative setting and obviously

does not prejudice the full welfare implications of the 2014 reform.

3.6 Conclusion

This paper studies the impact of property transaction taxes on a broad range of behavioral margins. Using exhaustive data on property transaction from 2010-2016, I investigate how housing markets reacted to a 2014 reform that let French counties raise stamp duties by 0.72pp. I find that tax increases were anticipated, salient and lead to significant re-timing responses from transaction partakers. However, I find no evidence of bunching at border tax notches suggesting that the tax increase was too small to outweigh buyers' location preferences.

Using a synthetic control approach pooling many low-level geographic units, I exploit the fact that the timing of reform adoption was phased-in across the country. I present DiD evidence that increasing taxes depressed market activity by 3.1%–3.5%, implying a tax elasticity in the range $[0.22, 0.25]$. However, it had no impact on prices nor quality of exchanged homes. I interpret these results in a matching framework with heterogeneous buyers and homogeneous homes. When taxation increases, buyers with lower home valuation decrease their search effort more, pushing up observed negotiated prices by letting higher-valuation buyers be matched more often. This improved matching efficiency may positively contribute to welfare.

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List of Figures

1.1	Distribution of Subsidy Stock Yearly Growth	19
1.2	Within-MF Public Good Changes	36
1.3	Within-MF Migration Response	37
1.4	Within-MF Per Capita Housing and Wage Responses	38
1.5	Between-MF Public Good Changes	40
1.6	Between-MF Migration Response	41
1.7	Between-MF Per Capita Housing and Wage Responses	42
1.8	Within-MF Housing Supply and Housing Price Responses	44
1.9	Between-MF Housing Supply and Housing Price Responses	44
1.10	Kernel Regressions: Within-MF Changes	49
1.11	Kernel Regressions: Between-MF Changes (1/2)	50
1.12	Kernel Regressions: Between-MF Changes (2/2)	51
2.1	Per Capita Spending vs. Population	72
2.2	Per Capita Spending vs. Density (Within Regressions)	74
2.3	Per Capita Spending vs. Density (Cross-Sectional Regressions)	75
2.4	Per Capita Spending vs. Effective Public Goods	77
2.5	Direction of Pareto-Improving Transfers ($\kappa = 0.44$)	96
2.6	Government's Beliefs	99
2.7	Distribution of Implied Weights for Various Planner's Beliefs	101
3.1	Treatment Groups	112
3.2	Reform Adoption Timeline	113
3.3	Summary Statistics	115
3.4	Bunching at the Time Notch	118

3.5	Bunching at the Border Notch	121
3.6	Gap in Number of Transactions	128
3.7	Gap in Normalized Number of Transactions	129
3.8	Gap in Cumulative Number of Transactions	130
3.9	Gap in Normalized m^2 Price	132
3.10	Quality Changes	133
3.11	Average Placebo Estimates	136
3.12	Treatment vs. Placebo Effects	137
A.1	Model Fit: G^f vs. G^s	164
A.2	Changes in (Public Employees) / (20–65 Population)	166
B.1	Sensibility Analysis for κ	185
B.2	Calibration of ψ	188
B.3	Skill Sorting	189

List of Tables

1.1	Descriptive Statistics on Jurisdictions	32
1.2	Within-MF Sensitivity Analysis	39
1.3	Between-MF Sensitivity Analysis	43
1.4	Housing Supply Sensitivity Analysis	45
1.5	Structural Parameters for $\alpha = 0.25$	55
1.6	Structural Parameters for $\alpha = 0.30$	55
1.7	Structural Parameters for $\alpha = 0.35$	56
1.8	Structural Parameters for $\alpha = 0.40$	56
1.9	Structural Parameters Calibration and Estimation	57
1.10	Welfare Change Estimates	60
2.1	Structure of Pareto Weights	102
3.1	Balancing Over-Identification Checks	126
3.2	Balancing Over-Identification Checks (Average Placebo)	135
B.1	Structural Elasticities for $\frac{1}{1+\rho_C} = 0.5$	184
B.2	Cross-Sectional Estimates	186
B.3	Linear Regression Share Estimates	186
B.4	Log-Linearity Test	187
B.5	Calibration of ψ	187

Appendix A

Appendix to “This Town Ain’t Big Enough? Quantifying Public Good Spillovers”

A.1 Jurisdiction Utility

The full expression for utility in jurisdiction j is

$$\begin{aligned} v_j = & (1 - \phi)(1 - \alpha) \ln(1 - \alpha) + (1 - \phi)(1 - \alpha) \ln(1 - \tau^w) + (1 - \phi)(1 - \alpha) \ln(w_j) \\ & + (1 - \phi)\alpha \ln(h_j) + \phi\delta \ln(G_j) + \phi(1 - \delta) \frac{1}{|a_j|} \sum_{j' \in a_j} \ln(G_{j'}) - \phi\delta^2 \kappa \ln(N_j) \\ & - 2\phi\delta(1 - \delta)\kappa \frac{1}{|a_j|} \sum_{j' \in a_j} \ln(N_{j'}) - \phi(1 - \delta)^2 \kappa \frac{1}{|a_j|} \sum_{j' \in a_j} \ln(N_{j'}) + \ln(\mathcal{E}_j^A) \end{aligned} \quad (\text{A.1})$$

A.2 Local Public Goods

Our baseline model features a simple representation of local public spending. In Appendix [A.2.1](#), we develop a more realistic version that we use in our empirical applications in Sections [1.5](#) and [1.6](#). Appendix [A.2.2](#) gives conditions under which our framework is time-consistent. We take our extended version to the data in Appendix [A.2.3](#).

A.2.1 Extension of the Model

We assume that public good quantity G produced by a jurisdiction has a Cobb-Douglas form:

$$\ln G = \varphi \ln G^s + (1 - \varphi) \ln G^f \quad (\text{A.2})$$

where G^s is the stock of public capital and G^f the flow of services annually consumed. Both are directly measured as spending in terms of the numéraire good and we abstract from differences in public good provision efficiency. Residents vote on the residence tax τ^h , the business tax τ^k and the amount of G^s and G^f . Because we assumed homogeneous preferences, the voting mechanism is akin to a maximization problem by a local social planner.

Residents commit to policy $\{G^s, G^f, \tau^h, \tau^k\}$ for current and all future periods.¹ They pay for G^f every year. Durable investments depreciate at the annual rate ρ . To maintain a constant level of infrastructure, residents hence have to pay G^s the first year and ρG^s every following year.

Cities inherit zero net wealth from the past.² They anticipate a flow of future subsidies $\{F_t\}_0^\infty$ and have access to international debt markets with fixed interest rate R . Residents' preferred policy is found by maximizing

$$(1 - \alpha) \ln \left((1 - \tau^w)w - r(1 + \tau^h)h \right) + \alpha \ln h + \frac{\phi}{1 - \phi} \delta \left(\varphi \ln G^s + (1 - \varphi) \ln G^f \right) \quad (\text{A.3})$$

over $\{G^s, G^f, \tau^h, \tau^k\}$ subject to jurisdiction inter-temporal budget constraint

$$\zeta G^s + G^f = \tau^h r H + \tau^k R K + \frac{R}{1 + R} \left[\sum_{t=0}^{\infty} \left(\frac{1}{1 + R} \right)^t F_t \right] \quad (\text{A.4})$$

where $\zeta = (\rho + R)/(1 + R)$.

A.2.2 Time Consistency and No Inherited Wealth

We show that jurisdictions' decisions are time consistent. We prove it in the case without local business tax, but the argument remains valid in the presence of business taxes. We assume that jurisdictions actually inherit investment G_{init}^s and debt D_{init} from the past. Residents' preferred policy is given by

$$\begin{aligned} \max_{G^s, G^f, \tau} & (1 - \alpha)(1 - \phi) \ln \left((1 - \tau^w)w - r(1 + \tau)h \right) + \alpha(1 - \phi) \ln(h) \\ & + \phi \left(\delta + \frac{1 - \delta}{|a_j|} \right) \left(\varphi \ln(G^s) + (1 - \varphi) \ln(G^f) \right) \end{aligned}$$

¹Appendix A.2.2 shows that policy choices are time-consistent if the environment stays constant. When there is a shock to the environment (e.g., in local amenities or public subsidies) jurisdictions change their equilibrium policy which is again time-consistent.

²Appendix A.2.2 shows that under no-Ponzi conditions, jurisdictions inherit exactly zero net wealth from past periods. Hence, there is no path-dependency in local public goods choices even in the presence of durable investments and a seemingly dynamic problem collapses into a static one.

That is, after substituting in optimal housing and numéraire consumption:

$$\max_{G^s, G^f, \tau} -\alpha(1 - \phi) \ln(1 + \tau) + \phi \left(\delta + \frac{1 - \delta}{|a_j|} \right) (\varphi \ln(G^s) + (1 - \varphi) \ln(G^f))$$

Optimal policy from $t=0$ (creation of the jurisdiction) In period $t = 0$ (e.g., the creation of the jurisdiction), jurisdictions inherit zero investments $G_{init}^s = 0$ and debt $D_{init} = 0$. Residents choose and commit to a constant level of public good G_0 and a constant tax rate τ_0 for current and all future periods $\{t = 0, t = 1, \dots\}$. They furthermore assume that population will not change in future periods. When population changes because of, say, an amenity shock, the problem is reinitialized at period 0, but this time with a history and an *a priori* non-zero initial jurisdiction net wealth. Public goods are durable but depreciate at speed ρ . To commit to the initially chosen G_0^s residents hence have to invest ρG_0^s every period from $t = 1$. Residents receive a flow of subsidy equal to F and constant over time (again, if F changes the problem is reinitialized at period 0 with an *a priori* non-zero initial net wealth). jurisdictions can levy debt each year $\{D_0, D_1, D_2, \dots\}$ that has to be repaid in full the next year, plus interests. Yearly budgets are given by

$$\begin{aligned} G_0^s + G_0^f &= D_0 + \tau_0 Nhr + F \\ \rho G_0^s + G_0^f &= -(1 + R)D_0 + D_1 + \tau_0 Nhr + F \\ \rho G_0^s + G_0^f &= -(1 + R)D_1 + D_2 + \tau_0 Nhr + F \\ &\vdots \\ \rho G_0^s + G_0^f &= -(1 + R)D_{t-1} + D_t + \tau_0 Nhr + F \\ &\vdots \end{aligned} \tag{A.5}$$

Let's multiply each budget line by

$$p_t \equiv \left(\frac{1}{1 + R} \right)^t \tag{A.6}$$

and sum all lines up to $t = T$:

$$G_0^s + \rho G_0^s \sum_{t=1}^T p_t + G_0^f \sum_{t=0}^T p_t = \sum_{t=0}^T p_t D_t - (1 + R) \sum_{t=1}^T p_t D_{t-1} + \tau_0 Nhr \sum_{t=0}^T p_t + F \sum_{t=0}^T p_t \tag{A.7}$$

That is:

$$G_0^s + G_0^f + (\rho G_0^s + G_0^f) \sum_{t=1}^T p_t = p_T D_T + \tau_0 Nhr \sum_{t=0}^T p_t + F \sum_{t=0}^T p_t \quad (\text{A.8})$$

We assume that the following no-Ponzi scheme condition holds:

$$\lim_{T \rightarrow \infty} p_T D_T = 0 \quad (\text{A.9})$$

and take the limit of (A.8) when $T \rightarrow \infty$ to get the inter-temporal budget constraint:

$$\zeta G_0^s + G_0^f = \tau_0 Nhr + F \quad (\text{A.10})$$

Note that we can express the debt stock at all t by solving

$$D_t = (1 + R)D_{t-1} + \rho G_0^s + G_0^f - \tau_0 Nhr - F = (1 + R)D_{t-1} + (\rho - 1)G_0^s + D_0$$

that is (noticing an arithmetico-geometric sequence):

$$D_t = (1 + R)^t \left(D_0 \frac{1 + R}{R} + (\rho - 1)G_0^s \frac{1}{R} \right) - \frac{1}{R} ((\rho - 1)G_0^s + D_0) \quad (\text{A.11})$$

That is:

$$p_t D_t = D_0 \frac{1 + R}{R} + (\rho - 1)G_0^s \frac{1}{R} - p_t \frac{1}{R} ((\rho - 1)G_0^s + D_0)$$

Taking the limit when $t \rightarrow \infty$ using the no-Ponzi condition yields

$$0 = D_0 \frac{1 + R}{R} + (\rho - 1)G_0^s \frac{1}{R} \quad (\text{A.12})$$

Plugging (A.12) in (A.11), we find that debt is constant and that we are in a stationary setting (holding environment fixed):

$$D_t = D_0 = \frac{1 - \rho}{1 + R} G_0 \quad (\text{A.13})$$

Optimal policy from $t=1$ (keeping environment fixed) At period 1, jurisdictions inherit debt (plus interests) and depreciated assets from period 0:

$$(1 - \rho)G_0^s - (1 + R)D_0$$

However, according to equation (A.12) this quantity is exactly zero. The optimization problem in period 1 is hence the same as in period 0, and policy choices from period 1 onward are unchanged compared to those in period 0. The proof by induction for any t follows straight-

forwardly.

Optimal policy from $t=1$ (with a change in the environment) Initial net jurisdiction wealth is zero so the problem is reinitialized at period zero following a shock to the environment. The environment has changed however, so policy choices may be different in the new equilibrium.

A.2.3 Taking the Extended Model to the Data

We construct the public good index G using our data on municipal accounts. First, we construct G^s and G^f , i.e., municipal durable facilities and annual public services. Then, we estimate parameter φ of the Cobb-Douglas function.

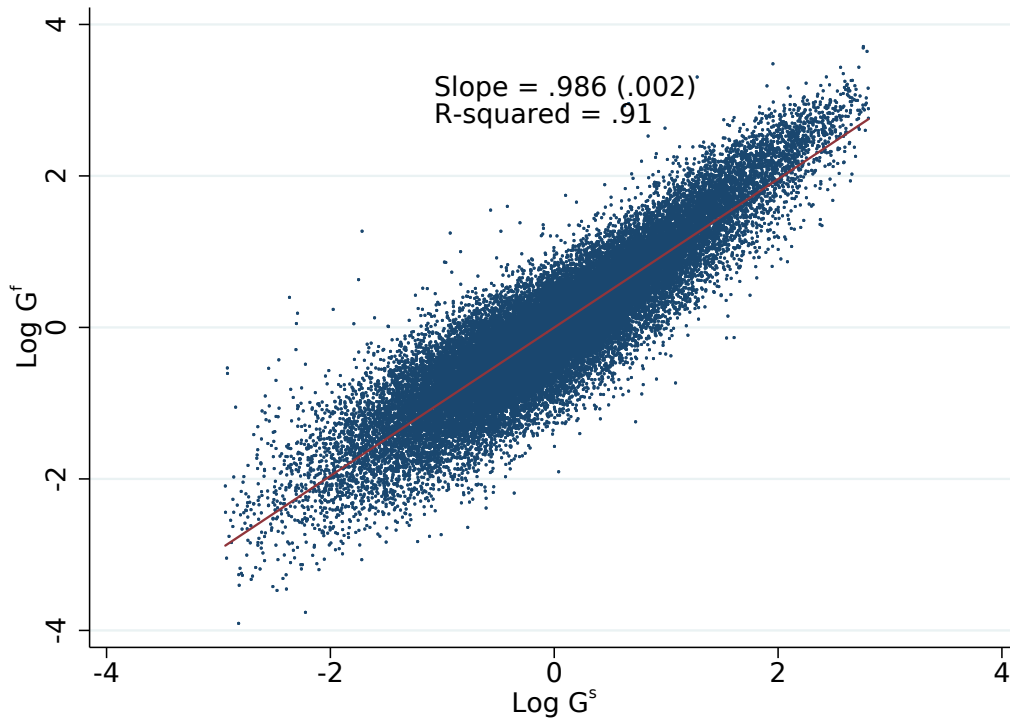
Construction of G^s We construct our public investment variable G^s as the sum of all public assets minus the raw value of the land and financial assets such as cash.³ They are recorded at book value and account for investment depreciation. As mentioned in Section 1.2, investments notably include schools, transportation infrastructure, parks improvements, sports facilities, museums, art collections, etc. They also include investment subsidies to other parties such as local clubs and associations. Importantly, G^s does not contain social housing units.⁴

Construction of G^f We construct G^f , the flow of services annually consumed, as the sum of staff expenditure, maintenance spending, payments for external services and operating subsidies to third parties. We exclude interests payments as they do not correspond to consumable services.

Construction of G To make sure that our constructed measures for G^s and G^f correctly reflect the amount of services chosen at the local level, we first check that they satisfy standard budget requirements. We then directly calibrate the ratio $\frac{G^s}{G^f}$ to recover parameter φ . We first estimate parameter ρ by calibrating asset depreciation based on municipal financial accounts. Our central estimate is $\rho = 0.010$. Then we calibrate parameter R from interests paid as a share of the debt stock. Our central estimate is $R = 0.041$. Both estimates are robust across

³Taking out the raw value of the land seems natural as residents are unlikely to value it. Given that land is mostly a gift from the central government, erasing this asset as well as the corresponding liability from the balance sheet is neutral in our analysis. Cash and other liquid assets can be considered negative debt and are accounted for in our theoretical framework in the form of future taxes.

⁴Social housing units, when publicly owned, are held by ad hoc entities and not by jurisdictions.

Figure A.1: Model Fit: G^f vs. G^s 

Note: This Figure shows the calibration exercise we carry out for our Cobb-Douglas modeling of how public goods enter utility.

different calibration methods. We are then able to calibrate parameter $\zeta = \frac{\rho+R}{1+R}$ with a central estimate of $\zeta = 0.049$. Finally, we estimate the ratio $\frac{G^s}{G^f} = \frac{\varphi}{(1-\varphi)\zeta}$ with a log-log regression reproduced in Figure A.1. The Figure comforts our Cobb-Douglas specification as the slope is close to unity and residuals are quite small as indicated by the large R^2 . Point estimate for the intercept is $1/0.074 \approx 13.5$, that is, capital investment approximately equals 13.5 years of operating expenditures. This leads to a central estimate $\varphi = 0.378$. We can then rewrite our public good index G as

$$G = G^s \left(\frac{1-\varphi}{\varphi} \zeta \right)^{1-\varphi} \approx 0.208 \times G^s$$

A.3 DiD Robustness Checks

Let jurisdiction total 20–65 population be

$$N = N_{\bar{p}} + N_p$$

where N_p is the public sector population. Let $s_p = N_p/N$ be the pre-shock share of public employees in the population. Call $\tilde{X} = \prod_{j' \in a_j} X_{j'}^{\frac{1}{|a_j|}}$. Within-MF changes in N following a shock can be decomposed as follows:

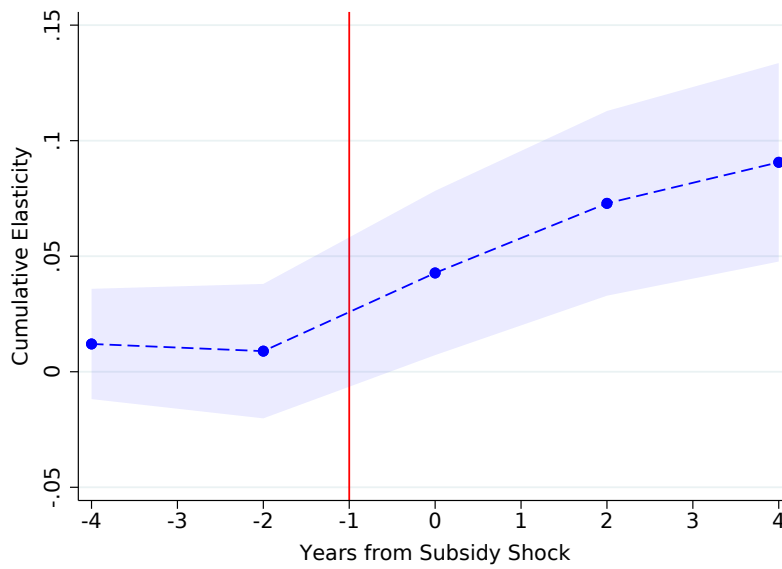
$$\begin{aligned} d \ln (N/\tilde{N}) &= \frac{d(N/\tilde{N})}{N/\tilde{N}} \\ &= \frac{d(N_{\bar{p}}/\tilde{N})}{N/\tilde{N}} + \frac{d(N_p/\tilde{N})}{N/\tilde{N}} \\ &= \frac{d(N_{\bar{p}}/\tilde{N})}{N/\tilde{N}} + \frac{d(s_p N/\tilde{N})}{N/\tilde{N}} \\ &= \frac{d(N_{\bar{p}}/\tilde{N})}{N/\tilde{N}} + s_p \left[\frac{d(N/\tilde{N})}{N/\tilde{N}} + \frac{ds_p}{s_p} \right] \end{aligned}$$

The share of migration responses coming from public employment is then

$$\frac{\frac{d(N_p/\tilde{N})}{N/\tilde{N}}}{\frac{d(N/\tilde{N})}{N/\tilde{N}}} = s_p \left[1 + \frac{ds_p/s_p}{\frac{d(N/\tilde{N})}{N/\tilde{N}}} \right]$$

Figure A.2 shows an estimate for $\frac{ds_p/s_p}{\frac{d(N/\tilde{N})}{N/\tilde{N}}}$ of around 0.09. The estimate for $\frac{\frac{d(N/\tilde{N})}{N/\tilde{N}}}{\frac{d(S/\tilde{S})}{S/\tilde{S}}}$ from Section 1.5.2 is around 0.03. The average estimate for the pre-shock ratio of public staff to total 20–65 population is $s_p \approx 3\%$. Estimate for the share of migration responses coming from public employment is hence $\approx 3\% \times (1 + 0.09/0.03) = 12\%$. We get similar estimates when including paid subsidies to third parties in addition to public staff payroll.

Figure A.2: Changes in (Public Employees) / (20–65 Population)



Note: This Figure plots the coefficient of regressions similar to Section 1.5 regressions. The dependent variable is the proportional change in the absolute share of public employees in municipality population $\Delta \ln(s_p)$ and the explanatory variable is the within-MF subsidy shock $\Delta \ln(\bar{S})$. Standard errors are clustered at the MF level. We report the 5% confidence bands.

A.4 Welfare

We detail here the calculations for the welfare analysis.

A.4.1 Before the Merger

We first back out the fundamentals of the model $\{\mathcal{E}_j^A, \mathcal{E}_j^C\}$ from (i) equilibrium endogenous variables of the model that we observe as econometricians in 2014, and (ii) parameters that we have estimated or calibrated from the literature. We then make several assumptions to ensure tractability of the equilibrium in our simplified setup. We assume away business taxes and subsidies from the central government and make five others assumptions: (i) residents are myopic and do not anticipate migration responses when they vote for public goods and taxes; (ii) the public good production function is the identity, i.e., $\Gamma(x) = x$; (iii) housing supply elasticities η_j are constant across jurisdictions; (iv) local productivity θ_j^Y does not depend on local public goods and (v) the geographic fundamentals that we back out from observed data are still defining geography in our simplified model. This version of the model has a unique and tractable equilibrium defined by the following equations:

$$N_j = \frac{\exp(v_j/\sigma)}{\sum_{j'} \exp(v_{j'}/\sigma)} \quad (\text{A.14})$$

$$\begin{aligned} v_j &= (1 - \phi)(1 - \alpha) \ln(1 - \alpha) + (1 - \phi)(1 - \alpha) \ln(w_j) \\ &+ (1 - \phi)\alpha \ln(h_j) + \phi\delta \ln(G_j) + \phi(1 - \delta) \frac{1}{|a_j|} \sum_{j' \in a_j} \ln(G_{j'}) - \phi\delta^2 \kappa \ln(N_j) \\ &- 2\phi\delta(1 - \delta)\kappa \frac{1}{|a_j|} \sum_{j' \in a_j} \ln(N_{j'}) - \phi(1 - \delta)^2 \kappa \frac{1}{|a_j|} \sum_{j' \in a_j} \ln(N_{j'}) + \ln(\mathcal{E}_j^A) \end{aligned} \quad (\text{A.15})$$

$$\ln(r_j) = \frac{1}{\eta_j} \ln\left(\frac{H_j}{T_j}\right) + \ln(\mathcal{E}_j^C) \quad (\text{A.16})$$

$$1 + \tau_j^h = 1 + \frac{\phi(\delta + \frac{1-\delta}{|a_j|})}{(1 - \phi)\alpha} \quad (\text{A.17})$$

$$G_j [1 + \tau_j^h] = \frac{\varphi^\varphi (1 - \varphi)^{1-\varphi}}{\zeta^\varphi} \tau_j^h N_j w_j \alpha \quad (\text{A.18})$$

Equations (A.17) and (A.18) are the first order conditions for public good supply in this simplified version.⁵ Equations (A.14), (A.15) and (A.16) are identical in the simplified model

⁵The vote conditions derive from the following optimization problem: $\max_{\tau_j^h, G_j^s, G_j^f, G_j} \left(G_j^\delta \prod_{j' \in a_j} G_{j'}^{\frac{1-\delta}{|a_j|}} \right)^\phi C_j^{1-\phi}$ under the constraints $\zeta G_j^s + G_j^f = N_j h_j r_j \tau_j^h$, $G_j = (G^s)^\varphi (G_j^f)^{1-\varphi}$,

and in our general model.

We note N^o , G^o , v^o , r^o and τ^o the equilibrium variables in the observed 2014 allocation and N^n , G^n , v^n , r^n and τ^n the pre-reform equilibrium variables in the simplified model that we show can be deduced from observed data.

We rewrite equilibrium conditions (A.14), (A.15) and (A.16):

$$N_j = \frac{\exp(v_j/\sigma)}{\sum_{j'} \exp(v_{j'}/\sigma)} \quad (\text{A.19})$$

$$v_j = -\frac{\eta}{\eta+1}(1-\phi)\alpha \ln(1+\tau_j) + \phi\delta \ln(G_j) + \phi \frac{1-\delta}{|a_j|} \sum_{j \in a_j} \ln G_{j'} \quad (\text{A.20})$$

$$\begin{aligned} & - \left[(1-\phi)\alpha \frac{1}{\eta+1} + \phi\kappa\delta^2 \right] \ln N_j - \phi\kappa \frac{1-\delta^2}{|a_j|} \sum_{j' \in a_j} \ln N_{j'} + \sigma\Theta_j^A \\ \ln r_j &= \frac{1}{\eta+1} \ln N_j - \frac{1}{\eta+1} \ln(1+\tau_j) + \Theta_j^r \end{aligned} \quad (\text{A.21})$$

with the transformed residuals

$$\begin{aligned} \Theta_j^r &= \frac{\eta}{\eta+1} \ln(\mathcal{E}_j^C) + \frac{1}{\eta+1} \ln \frac{\alpha w_j}{T_j} \\ \sigma\Theta_j^A &= (1-\phi)\alpha \ln \alpha + (1-\phi)(1-\alpha) \ln(1-\alpha) + (1-\phi) \ln(w_j) - (1-\phi)\alpha\Theta_j^r + \phi\delta\Theta_j^G \\ &+ \phi \frac{1-\delta}{|a_j|} \sum_{j \in a_j} \Theta_{j'}^G + \ln(\mathcal{E}_j^A) \end{aligned}$$

Let's define the following operators:

- $\underline{X}_j = X_j - \sum_{j \in a} \frac{X_j}{|a_j|}$
- $\bar{X}_j = \exp(\underline{\ln}(X_j))$
- $\underline{X}_{a_j} = \sum_{j' \in a} \frac{X_{j'}}{|a|} - \sum_{a'} \sum_{j' \in a'} \frac{X_{j'}}{J|a'|}$
- $\widehat{X}_j = \exp(\widehat{\ln}(X_j))$
- $\hat{X}_j = X_j - \sum_{j'} \frac{X_{j'}}{J}$
- $\check{X}_j = \exp(\hat{\ln}(X_j))$

$C_j = c_j^{1-\alpha} h_j^\alpha$, $c_j = (1-\alpha)w_j$ and $h_j = \frac{\alpha w_j}{r_j(1+\tau_j^h)}$. In the main text, $\Phi = \frac{\varphi^\varphi(1-\varphi)^{1-\varphi}}{\zeta^\varphi}$.

Applying the \underline{L} operator to equations (A.17), (A.18), (A.19), and (A.20) gives:

$$\left[\sigma + (1 - \phi)\alpha \frac{1}{\eta + 1} + \phi\kappa\delta^2 - \phi\delta \right] \ln \frac{\overline{N}_j^n}{\overline{N}_j^o} = \phi\delta \left(\ln \overline{N}_j^o - \ln \overline{G}_j^o + \ln \overline{w}_j \right) \quad (\text{A.22})$$

$$+ \frac{\eta}{\eta + 1} (1 - \phi)\alpha \ln \overline{1 + \tau}_j^o \quad (\text{A.23})$$

Applying the \underline{L} operator to equations (A.18), (A.19), and (A.20) we get:

$$\left[\sigma + \frac{(1 - \phi)\alpha}{(\eta + 1)} + \phi\kappa - \phi \right] \ln \frac{\widehat{N}_a^n}{\widehat{N}_a^o} = \phi \left(\ln \widehat{N}_a^o - \ln \widehat{G}_a^o + \ln \widehat{w}_a + \ln \widehat{\tau}_a^n - \ln \widehat{1 + \tau}_a^n \right) \quad (\text{A.24})$$

$$- \frac{\eta}{\eta + 1} (1 - \phi)\alpha \ln \frac{\widehat{1 + \tau}_a^n}{\widehat{1 + \tau}_a^o}$$

Notice that we can write:

$$\frac{N_j^n}{N_j^o} = \frac{\widehat{N}_j^n \overline{N}_j^n \sum_{j'} \widehat{N}_{j'}^o \overline{N}_{j'}^o}{\widehat{N}_j^o \overline{N}_j^o \sum_{j'} \widehat{N}_{j'}^n \overline{N}_{j'}^n} \quad (\text{A.25})$$

With equations (A.17), (A.23), (A.24), (A.25) and (A.18)—in this order—we get τ_j^n , N_j^n and G_j^n .

A.4.2 After the merger

We now assume that all jurisdictions within a municipal federation cooperate for the supply of public goods. Following the merger, residents vote for a new level of public goods pooling resources of all former member jurisdictions. To keep welfare comparable between pre- and post-reform situations, we assume that pre-reform jurisdictions still exist for the purpose of labor and housing markets. The only thing that changes is the level at which public good is supplied.

Housing tax rates are fixed by the MF council whose preferences exactly reflect those of all inhabitants in the MF. For a given MF a the council chooses the housing tax rate and the level of local public good G which maximize v_j —with j a jurisdiction belonging to the MF—the budget constraint $\sum_{j' \in a} r_{j'} N_{j'} h_{j'} \tau_{j'}^h = G$. Importantly, the optimization problem gives the same G and τ^h for all j belonging to a . Once G is chosen, a share is allocated to each jurisdiction. We assume that the allocation rule keeps public good shares unchanged, that is, jurisdiction j gets a share $\frac{G_j^n}{\sum_{j' \in a_j} G_{j'}^n}$ of G .

We note N^m , G^m , v^m , r^m and τ^m the equilibrium variable in the post-reform case. The new vote conditions are⁶

$$1 + \tau_a^m = 1 + \frac{\phi}{(1 - \phi)\alpha} \quad (\text{A.26})$$

$$G_j^m = \frac{\varphi^\varphi(1 - \varphi)^{1-\varphi}}{\zeta^\varphi} \frac{G_j^n}{\sum_{j' \in a_j} G_{j'}^n} \sum_{j' \in a_j} \frac{\tau_{j'}^m}{1 + \tau_{j'}^m} N_{j'}^m w_{j'} \alpha \quad (\text{A.27})$$

where τ_a^m and G_a^m are independent of the allocation rule chosen.

Equations (A.19), (A.20) and (A.21) still hold. We now rewrite N^m , G^m , v^m , r^m and τ^m as function of N^n , G^n , v^n , r^n and τ^n .

Applying the \underline{L} operator to equations (A.17), (A.18), (A.19), (A.20), (A.26), and (A.27) we get:

$$\bar{N}_j^m = \bar{N}_j^n \quad (\text{A.28})$$

that is, population shares within any MF stays constant. Note that

$$\begin{aligned} G_j^m &= \frac{G_j^n}{\sum_{j' \in a_j} G_{j'}^n} \sum_{j' \in a} G_{j'}^m \frac{\tau_{j'}^m}{1 + \tau_{j'}^m} \frac{1 + \tau_{j'}^n}{\tau_{j'}^n} \frac{N_{j'}^m}{N_{j'}^n} \\ &= \left[\prod_{j' \in a_j} \frac{N_{j'}^m}{N_{j'}^n} \right]^{\frac{1}{|a_j|}} \frac{G_j^n}{\sum_{j' \in a_j} G_{j'}^n} \sum_{j' \in a_j} G_{j'}^m \frac{\tau_{j'}^m}{1 + \tau_{j'}^m} \frac{1 + \tau_{j'}^n}{\tau_{j'}^n} \frac{\bar{N}_{j'}^m}{\bar{N}_{j'}^n} \\ &= \left[\prod_{j' \in a_j} \frac{N_{j'}^m}{N_{j'}^n} \right]^{\frac{1}{|a_j|}} \frac{G_j^n}{\sum_{j' \in a_j} G_{j'}^n} \sum_{j' \in a_j} G_{j'}^m \frac{\tau_{j'}^m}{1 + \tau_{j'}^m} \frac{1 + \tau_{j'}^n}{\tau_{j'}^n} \end{aligned} \quad (\text{A.29})$$

We define X_a such as

$$\frac{X_a}{G_j^n} = \frac{\sum_{j' \in a} G_{j'}^n \frac{\tau_{j'}^m}{1 + \tau_{j'}^m} \frac{1 + \tau_{j'}^n}{\tau_{j'}^n}}{\sum_{j' \in a_j} G_{j'}^n} \quad (\text{A.30})$$

Applying the \underline{L} operator to equations (A.18), (A.19), (A.20), (A.26), (A.27), (A.29) and (A.30)

⁶The vote conditions derive from the following optimization problem: $\max_{\tau_{a_j}^H, G_{a_j}, G_{a_j}^s, G_{a_j}^f} \left([G_{a_j} w_j]^\delta \prod_{j' \in a_j} [G_{a_j} w_{j'}]^{\frac{1-\delta}{|a_j|}} \right)^\phi C_j^{1-\phi}$ under the constraints $\zeta G_{a_j}^s + G_{a_j}^f = \sum_{j' \in a_j} N_j h_j r_j \tau_{a_j}^H$, $G_{a_j} = (G_{a_j}^s)^\varphi (G_{a_j}^f)^{1-\varphi}$, $C_j = c_j^{1-\alpha} h_j^\alpha$, $c_j = (1 - \alpha) w_j$, $h_j = \frac{\alpha w}{r_j(1 + \tau_{a_j}^H)}$.

we get:

$$\left[\sigma + (1 - \phi)\alpha \frac{1}{\eta + 1} + \phi\kappa - \phi \right] \ln \frac{\widehat{N}_j^m}{\widehat{N}_j^n} = \phi \ln \frac{\widehat{X}_a}{\widehat{G}_a^n} - \frac{\eta}{\eta + 1} (1 - \phi)\alpha \ln \frac{\widehat{1 + \tau_a^m}}{\widehat{1 + \tau_a^n}} \quad (\text{A.31})$$

Note that we have

$$\frac{N_j^m}{N_j^n} = \frac{\widehat{N}_j^m \overline{N}_j^m}{\widehat{N}_j^n \overline{N}_j^n} \frac{\sum_{j'} \widehat{N}_{j'}^n \overline{N}_{j'}^n}{\sum_{j'} \widehat{N}_{j'}^m \overline{N}_{j'}^m} \quad (\text{A.32})$$

With equations (A.26), (A.28), (A.29), (A.30), (A.31), (A.32) and (A.27) — in this order — we get τ_j^m , N_j^m and G_j^m .

Finally:

$$\begin{aligned} v_j^m - v_j^n &= \phi\delta \ln \frac{G_j^m}{G_j^n} + \phi \frac{1 - \delta}{|a_j|} \sum_{j \in a_j} \ln \frac{G_{j'}^m}{G_{j'}^n} - \left[(1 - \phi)\alpha \frac{1}{\eta + 1} + \phi\kappa\delta^2 \right] \ln \frac{N_j^m}{N_j^n} \\ &\quad - \phi\kappa \frac{1 - \delta^2}{|a_j|} \sum_{j' \in a_j} \ln \frac{N_{j'}^m}{N_{j'}^n} - \frac{\eta}{\eta + 1} (1 - \phi)\alpha \ln \left(\frac{1 + \tau_{a_j}^m}{1 + \tau_{a_j}^n} \right) \end{aligned} \quad (\text{A.33})$$

A.4.3 Welfare Comparison

We define social welfare as

$$W = \mathbb{E} \left[\max_j (v_j + \mu_{ij}) \right]$$

Since idiosyncratic tastes are distributed Extreme Value Type-I, we have⁷

$$W = \sigma\zeta + \sigma \ln \left(\sum_j \exp(v_j/\sigma) \right)$$

hence

$$W^o = \sigma\zeta + \sum_j \frac{v_j^o}{J} + \sigma \ln \left(\sum_j \exp \left(\frac{v_j - \sum_j \frac{v_j^o}{J}}{\sigma} \right) \right)$$

⁷ ζ is the Euler constant.

Note that $\check{N}_j^n = \exp\left(\frac{\widehat{v}_j^n}{\sigma}\right)$ therefore

$$W^n = \sigma\zeta + \sum_j \frac{v_j^n}{J} + \sigma \ln \left(\sum_j \check{N}_j^n \right)$$

and

$$W^m = \sigma\zeta + \sum_j \frac{v_j^n}{J} + \sigma \ln \left[\sum_j \check{N}_j^n \exp \frac{v_j^m - v_j^n}{\sigma} \right]$$

Eventually

$$W^m - W^n = \sigma \ln \left(\frac{\sum_j \check{N}_j^n \exp \frac{v_j^m - v_j^n}{\sigma}}{\sum_j \check{N}_j^n} \right) \quad (\text{A.34})$$

A.4.4 Welfare Estimation

For the estimation of welfare change, we use parameters estimates of Section 1.6 and their estimated variance-covariance structure. As we have as many parameters estimates as we have moment specifications, we take the average across specifications for each parameter. We compute Monte Carlo standard errors by simulating 10,000 new values for each parameter using the estimated variance-covariance matrix. We then estimate the welfare impact of a reform for each of these simulated set of parameters and compute percentile-based standard errors. In our baseline scenario, we take $\eta = 0.2$ and $\alpha = 0.3$.

Appendix B

Appendix to “Optimal Spatial Policies with Public Goods and Unobserved Location Preferences”

B.1 Local Public Good Demand

We generalize the arguments in [Boadway \(1982\)](#) who separately studies residence and property taxes in a two-city and homogeneous location preference model to our framework with many cities $J > 2$, heterogeneous preferences and simultaneous tax instruments. First, note that (2.10) and (2.11) can be combined to solve for $h_j(w_j, r_j, T_j, G_j, \tau_j^h, \{\Pi_k\}_k, \{\tau_k^p\}_k)$. Normalizing all prices with respect to p and with a slight abuse of notations, the local government

$$\max v_j \left(w_j + \sum_k \Pi_k (1 - \tau_k^p) + T_j - \tau_j^h - r_j h_j, h_j, \mathbf{G}_j(G_j, N_j), \mathcal{G} \right) \quad (\text{B.1})$$

over $\{\tau_j^h, \tau_j^p, G_j\}$ subject to

$$\Pi_j = F_j^Y + r_j F_j^H + p_j^G F_j^G - w_j (L_j^Y + L_j^H + L_j^G) \quad (\text{B.2})$$

$$N_j = \frac{v_j^{1/\sigma}}{\sum_k v_k^{1/\sigma}} \quad (\text{B.3})$$

$$L_j^Y + L_j^G + L_j^H + L_j^O = N_j \quad (\text{B.4})$$

$$F_j^H(L_j^H) = N_j h_j \quad (\text{B.5})$$

$$F_j^G(L_j^G) = G_j \quad (\text{B.6})$$

$$w_j = \frac{dF_j^Y(L_j^Y)}{dL_j^Y} \quad (\text{B.7})$$

$$w_j = r_j \frac{dF_j^H(L_j^H)}{dL_j^H} \quad (\text{B.8})$$

$$w_j = p_j^G \frac{dF_j^G(L_j^G)}{dL_j^G} \quad (\text{B.9})$$

$$p_j^G G_j = \tau_j^h N_j + \tau_j^p \Pi_j \quad (\text{B.10})$$

where policy instruments of other jurisdictions are considered fixed by the local government (Cournot-Nash equilibrium).

Plugging in as many constraints as possible, one can show that the first-order condition for τ_j^h yields

$$\underbrace{\frac{N_j}{p_j^G} \frac{\partial v_j}{\partial \mathbf{G}_j} \frac{\partial \mathbf{G}_j}{\partial G_j} - \frac{\partial v_j}{\partial c_j}}_{\text{Myopia Term}} + \underbrace{\frac{v_j}{N_j} \frac{\partial \ln v_j}{\partial \ln N_j} \Big|_{\tau} \frac{\partial N_j}{\partial \tau_j^h}}_{\text{Anticipation Term}} = 0 \quad (\text{B.11})$$

where $\frac{\partial \ln v_j}{\partial \ln N_j} \Big|_{\tau}$ is the change in local utility only coming from migration responses — affecting the tax base, local labor and goods markets, profits and public good congestion — holding constant all policy instruments. The term $\frac{\partial N_j}{\partial \tau_j^h}$ is the total or “policy” variation of N_j with respect to a change in τ_j^h when all other policy instruments stay constant.

The first two terms on the left-hand side correspond to the optimization of a myopic local government. One recognizes the Samuelson rule generalized to accommodate congestion. The last term corresponds to the migration anticipation term for a non-myopic local planner. Using (B.3), one can further show that

$$\frac{\partial N_j}{\partial \tau_j^h} \left(\sigma - \frac{\partial \ln v_j}{\partial \ln N_j} \right) = N_j (1 - N_j) \frac{1}{v_j} \left(\frac{N_j}{p_j^G} \frac{\partial v_j}{\partial \mathbf{G}_j} \frac{\partial \mathbf{G}_j}{\partial G_j} - \frac{\partial v_j}{\partial c_j} \right) - N_j \sum_l \frac{\partial \ln v_l}{\partial \ln N_l} \frac{\partial N_l}{\partial \tau_j^h} \quad (\text{B.12})$$

and that

$$\frac{\partial N_k}{\partial \tau_j^h} \left(\sigma - \frac{\partial \ln v_k}{\partial \ln N_k} \right) = -N_k N_j \frac{1}{v_j} \left(\frac{N_j}{p_j^G} \frac{\partial v_j}{\partial \mathbf{G}_j} \frac{\partial \mathbf{G}_j}{\partial G_j} - \frac{\partial v_j}{\partial c_j} \right) - N_k \sum_l \frac{\partial \ln v_l}{\partial \ln N_l} \frac{\partial N_l}{\partial \tau_j^h} \quad \text{for } k \neq j \quad (\text{B.13})$$

Plugging the last term of (B.12) in (B.13) one gets

$$\frac{\partial N_k}{\partial \tau_j^h} = \frac{\partial N_j}{\partial \tau_j^h} \frac{N_k}{N_j} \frac{\sigma - \frac{\partial \ln v_j}{\partial \ln N_j}}{\sigma - \frac{\partial \ln v_k}{\partial \ln N_k}} - \left(\frac{N_j}{p_j^G} \frac{\partial v_j}{\partial \mathbf{G}_j} \frac{\partial \mathbf{G}_j}{\partial G_j} - \frac{\partial v_j}{\partial c_j} \right) \frac{N_k}{v_j} \frac{1}{\sigma - \frac{\partial \ln v_k}{\partial \ln N_k}} \quad \text{for } k \neq j \quad (\text{B.14})$$

Plugging (B.14) in (B.12) for each $l \neq j$ we get

$$\frac{\partial N_j}{\partial \tau_j^h} = \frac{N_j}{v_j} A_j \times \left(\frac{N_j}{p_j^G} \frac{\partial v_j}{\partial \mathbf{G}_j} \frac{\partial \mathbf{G}_j}{\partial G_j} - \frac{\partial v_j}{\partial c_j} \right) \quad (\text{B.15})$$

where

$$A_j = \frac{1 - N_j + \sum_k N_k \frac{\partial \ln v_k}{\partial \ln N_k} \left(\sigma - \frac{\partial \ln v_k}{\partial \ln N_k} \right)^{-1}}{\sigma - \frac{\partial \ln v_j}{\partial \ln N_j} + \sum_k N_k \frac{\partial \ln v_k}{\partial \ln N_k} \left(\sigma - \frac{\partial \ln v_j}{\partial \ln N_j} \right) \left(\sigma - \frac{\partial \ln v_k}{\partial \ln N_k} \right)^{-1}}$$

so that finally first-order conditions for τ_j^h can be expressed as

$$\left(\frac{N_j}{p_j^G} \frac{\partial v_j}{\partial \mathbf{G}_j} \frac{\partial \mathbf{G}_j}{\partial G_j} - \frac{\partial v_j}{\partial c_j} \right) \left(1 + \frac{\partial \ln v_j}{\partial \ln N_j} \Big|_{\tau} \times A_j \right) = 0 \quad (\text{B.16})$$

and the only way to nullify (B.16) is to satisfy a local Samuelson rule.

The first-order condition for τ_j^p yields

$$\underbrace{\Pi_j \left(\frac{1}{p_j^G} \frac{\partial v_j}{\partial \mathbf{G}_j} \frac{\partial \mathbf{G}_j}{\partial G_j} - \frac{\partial v_j}{\partial c_j} \right)}_{\text{Myopia Term}} + \underbrace{\frac{v_j}{N_j} \frac{\partial \ln v_j}{\partial \ln N_j} \Big|_{\tau} \frac{\partial N_j}{\partial \tau_j^p}}_{\text{Anticipation Term}} = 0 \quad (\text{B.17})$$

One can see that (B.11) and (B.17) cannot hold simultaneously at interior solutions for τ_j^h and τ_j^p if the local planner is myopic about migration. Hence, the only way to have interior solutions for both head taxes and profit taxes is to have a non-myopic local planner. Using (B.3), one can further show that

$$\begin{aligned} \frac{\partial N_j}{\partial \tau_j^p} \left(\sigma - \frac{\partial \ln v_j}{\partial \ln N_j} \right) &= \Pi_j N_j (1 - N_j) \frac{1}{v_j} \left(\frac{1}{p_j^G} \frac{\partial v_j}{\partial \mathbf{G}_j} \frac{\partial \mathbf{G}_j}{\partial G_j} - \frac{\partial v_j}{\partial c_j} \right) \\ &+ N_j \sum_{l \neq j} R_l(N_l) \frac{N_l}{v_l} \frac{\partial v_l}{\partial c_l} - N_j \sum_l \frac{\partial \ln v_l}{\partial \ln N_l} \frac{\partial N_l}{\partial \tau_j^p} \end{aligned} \quad (\text{B.18})$$

$$\begin{aligned} \frac{\partial N_k}{\partial \tau_j^p} \left(\sigma - \frac{\partial \ln v_k}{\partial \ln N_k} \right) &= -\Pi_k N_k \frac{1}{v_k} \frac{\partial v_k}{\partial c_k} - N_j N_j \Pi_j \frac{1}{v_j} \left(\frac{1}{p_j^G} \frac{\partial v_j}{\partial \mathbf{G}_j} \frac{\partial \mathbf{G}_j}{\partial G_j} - \frac{\partial v_j}{\partial c_j} \right) \\ &\quad + N_k \sum_{l \neq j} R_l(N_l) \frac{N_l}{v_l} \frac{\partial v_l}{\partial c_l} - N_k \sum_l \frac{\partial \ln v_l}{\partial \ln N_l} \frac{\partial N_l}{\partial \tau_j^p} \end{aligned} \quad (\text{B.19})$$

$$\begin{aligned} \frac{\partial N_k}{\partial \tau_j^p} &= \frac{\partial N_j}{\partial \tau_j^p} \frac{N_k}{N_j} \frac{\sigma - \frac{\partial \ln v_j}{\partial \ln N_j}}{\sigma - \frac{\partial \ln v_k}{\partial \ln N_k}} - \Pi_k N_k \frac{1}{v_k} \frac{\partial v_k}{\partial c_k} \frac{1}{\sigma - \frac{\partial \ln v_k}{\partial \ln N_k}} \\ &\quad - \Pi_j \frac{1}{v_j} \left(\frac{1}{p_j^G} \frac{\partial v_j}{\partial \mathbf{G}_j} \frac{\partial \mathbf{G}_j}{\partial G_j} - \frac{\partial v_j}{\partial c_j} \right) \frac{(1 - N_j) N_k + N_j N_j}{\sigma - \frac{\partial \ln v_k}{\partial \ln N_k}} \quad \text{for } k \neq j \end{aligned} \quad (\text{B.20})$$

so

$$\Pi_j \left(\frac{1}{p_j^G} \frac{\partial v_j}{\partial \mathbf{G}_j} \frac{\partial \mathbf{G}_j}{\partial G_j} - \frac{\partial v_j}{\partial c_j} \right) \left(1 + \frac{\partial \ln v_j}{\partial \ln N_j} \Big|_{\tau} \times B_j \right) + C_j \sum_{k \neq j} \Pi_k N_k \frac{v_j}{v_k} \frac{\partial v_k}{\partial c_k} = 0 \quad (\text{B.21})$$

where

$$\begin{aligned} B_j &= \frac{1 - N_j + \sum_k N_k \frac{\partial \ln v_k}{\partial \ln N_k} \left(\sigma - \frac{\partial \ln v_k}{\partial \ln N_k} \right)^{-1} (N_k(1 - N_j) + N_j N_j)}{\sigma - \frac{\partial \ln v_j}{\partial \ln N_j} + \sum_k N_k \frac{\partial \ln v_k}{\partial \ln N_k} \left(\sigma - \frac{\partial \ln v_j}{\partial \ln N_j} \right) \left(\sigma - \frac{\partial \ln v_k}{\partial \ln N_k} \right)^{-1}} \\ C_j &= \frac{1}{\sigma - \frac{\partial \ln v_j}{\partial \ln N_j} + \sum_k N_k \frac{\partial \ln v_k}{\partial \ln N_k} \left(\sigma - \frac{\partial \ln v_j}{\partial \ln N_j} \right) \left(\sigma - \frac{\partial \ln v_k}{\partial \ln N_k} \right)^{-1}} \end{aligned}$$

The existence of the last term in the left-hand side of (B.21) ensures that the local Samuelson rule implied by (B.16) can be respected as an interior condition for τ_j^h . Absent this last term, the first term in large brackets in (B.21) would have to be zero which would be incompatible with the Samuelson rule.

B.2 A Two-Region Example

Traded and local public goods are produced using CRS technologies with exogenous local productivity $z_j = z_j^Y = z_j^G$ which is constant in both sectors and varies between cities with $z_1 > z_2$. This is the only source of spatial heterogeneity. Labor demand optimization yields

$$w_1 = z_1 > w_2 = z_2$$

Using (2.12) and (2.14), it implies that $p_j^G = p$ for all j . Agent i living in j inelastically supplies one unit of labor. She consumes the nationally traded good, housing and has utility

$$v_{ij} = v_j \mu_{ij}$$

where

$$v_j = \left(c_j^{1-\alpha} h_j^\alpha \right)^{1-\phi} (\mathbf{G})^\phi$$

Endogenous public goods $\mathbf{G} = \frac{1}{p} \tau_j^h N_j^{1-\kappa}$ are financed by the residence tax τ_j^h only and increase with population N_j with constant agglomeration elasticity $1 - \kappa > 0$. Normalizing $p = 1$, household budget is

$$z_j + T_j + \Pi = c_j + r_j h_j + \tau_j^h$$

where $\Pi = r_1 (L_1^H)^{\frac{\eta}{1+\eta}} + r_2 (L_2^H)^{\frac{\eta}{1+\eta}} - w_1 L_1^H - w_2 L_2^H$ are profits from the housing sector. Demand for traded and non-traded goods is $c_j = (1 - \alpha) (1 - \phi) (w_j + T_j + \Pi)$ and $r_j h_j = \alpha (1 - \phi) (w_j + T_j + \Pi)$ while public good demand is $\tau_j^h = \phi (w_j + T_j + \Pi)$. Combining these expressions yields

$$v_j = \frac{w_j + T_j + \Pi}{r_j^{\alpha(1-\phi)}} N_j^{(1-\kappa)\phi}$$

omitting multiplicative constants. Assuming constant housing supply elasticity η , we combine housing supply $r_j = w_j \frac{\eta+1}{\eta} h^{\frac{1}{\eta}} N^{\frac{1}{\eta}}$ with housing demand to get

$$r_j^{1+\eta} = \left(\frac{\eta+1}{\eta} \right)^\eta w_j^\eta N_j^\eta \alpha (1 - \phi) (w_j + T_j + \Pi)$$

which finally yields

$$v_j = (w_j + T_j + \Pi)^{1 - \frac{\alpha(1-\phi)}{1+\eta}} N_j^{(1-\kappa)\phi - \frac{\alpha(1-\phi)}{1+\eta}}$$

omitting multiplicative constants again.

The μ_{ij} 's are distributed Extreme Value Type-I $\{0, \sigma\}$ with dispersion $\sigma > 0$. Demand for region j is

$$N_j = \frac{v_j^{1/\sigma}}{v_1^{1/\sigma} + v_2^{1/\sigma}}$$

The central government sets up a per capita tax and transfer scheme $\{T_1, T_2\}$ where T_1 is chosen freely and T_2 endogenously adjusts to balance the budget, i.e., $N_1 T_1 + N_2 T_2 = 0$.

Equilibrium An equilibrium given T_1 is characterized by T_2 , Π and $\{N_j, L_j^Y, L_j^H, L_j^G, T_j, r_j\}_j$ such that

$$\begin{aligned}
\frac{\sigma - (1 - \kappa)\phi + \frac{\alpha(1-\phi)}{1+\eta}}{1 - \frac{\alpha(1-\phi)}{1+\eta}} \ln\left(\frac{N_1}{N_2}\right) &= \ln\left(\frac{z_1 + T_1 + \Pi}{z_2 + T_2 + \Pi}\right) \\
N_1 + N_2 &= 1 \\
N_1 T_1 + N_2 T_2 &= 0 \\
L_1^Y + L_1^H + L_1^G &= N_1 \\
L_2^Y + L_2^H + L_2^G &= N_2 \\
N_1 \frac{1}{r_1} \alpha(1 - \phi)(z_1 + T_1 + \Pi) &= (L_1^H)^{\eta/(1+\eta)} \\
N_2 \frac{1}{r_2} \alpha(1 - \phi)(z_2 + T_2 + \Pi) &= (L_2^H)^{\eta/(1+\eta)} \\
\phi(z_1 + T_1 + \Pi) N_1^{1-\kappa} &= z_1 L_1^G \\
\phi(z_2 + T_2 + \Pi) N_2^{1-\kappa} &= z_2 L_2^G \\
\Pi &= r_1 (L_1^H)^{\frac{\eta}{1+\eta}} + r_2 (L_2^H)^{\frac{\eta}{1+\eta}} - z_1 L_1^H - z_2 L_2^H \\
r_1^{1+\eta} &= \left(\frac{\eta+1}{\eta}\right)^\eta z_1^\eta N_1 \alpha(1 - \phi)(z_1 + T_1 + \Pi) \\
r_2^{1+\eta} &= \left(\frac{\eta+1}{\eta}\right)^\eta z_2^\eta N_2 \alpha(1 - \phi)(z_2 + T_2 + \Pi)
\end{aligned} \tag{B.22}$$

where we impose that dispersion forces — land congestion and location preferences — are stronger than agglomeration forces:

$$\sigma + \frac{\alpha(1 - \phi)}{1 + \eta} > (1 - \kappa)\phi$$

to ensure that the equilibrium is unique, stable and that the planner’s objective is concave. Walras’ law ensures that the market for the traded good also clears:

$$z_1 L_1^Y + z_2 L_2^Y = N_1(1 - \alpha)(1 - \phi)(z_1 + T_1 + \Pi) + N_2(1 - \alpha)(1 - \phi)(z_2 + T_2 + \Pi)$$

Small reform in the laissez-faire economy Let us introduce a small transfer dT_1 in region 1 from $T_1 = T_2 = 0$ while T_2 adjusts endogenously to balance the central budget. Implicitly differentiating the first three equilibrium equations of (B.22), population grows in region 1 according to:

$$\frac{\sigma - (1 - \kappa)\phi + \frac{\alpha(1-\phi)}{1+\eta}}{1 - \frac{\alpha(1-\phi)}{1+\eta}} \frac{dN_1}{dT_1} = N_1 \left(\frac{N_2}{z_1 + \Pi} + \frac{N_1}{z_2 + \Pi} \right) + \frac{d\Pi}{dT_1} N_2 N_1 \left(\frac{1}{z_1 + \Pi} - \frac{1}{z_2 + \Pi} \right)$$

and, using profit maximization from the three production sectors, profits change according to:

$$\frac{d\Pi}{dT_1} = \frac{\frac{\alpha(1-\phi)}{1+\eta}}{1 - \frac{\alpha(1-\phi)}{1+\eta}} (z_1 - z_2) \frac{dN_1}{dT_1} \quad (\text{B.23})$$

so that

$$\frac{dN_1}{dT_1} = \frac{N_1 \left(\frac{N_2}{z_1 + \Pi} + \frac{N_1}{z_2 + \Pi} \right)}{\frac{\sigma - (1-\kappa)\phi + \frac{\alpha(1-\phi)}{1+\eta}}{1 - \frac{\alpha(1-\phi)}{1+\eta}} - N_2 N_1 \left(\frac{1}{z_1 + \Pi} - \frac{1}{z_2 + \Pi} \right) \frac{\frac{\alpha(1-\phi)}{1+\eta}}{1 - \frac{\alpha(1-\phi)}{1+\eta}} (z_1 - z_2)} > 0 \quad (\text{B.24})$$

since $z_1 > z_2$. Note that it implies $\frac{d\Pi}{dT_1} > 0$.

Using a standard envelope argument, welfare of households marginally indifferent between 1 and 2 is not affected to a first order. Notice that the μ_{ij} 's of infra-marginals are unaffected. Log utility of infra-marginal residents in both cities changes as follows:

$$\begin{aligned} \frac{d \ln v_1}{dT_1} &\propto \frac{1}{z_1 + \Pi} + \frac{1}{z_1 + \Pi} \frac{d\Pi}{dT_1} + \frac{(1-\kappa)\phi - \frac{\alpha(1-\phi)}{1+\eta}}{1 - \frac{\alpha(1-\phi)}{1+\eta}} \frac{1}{N_1} \frac{dN_1}{dT_1} \\ \frac{d \ln v_2}{dT_1} &\propto -\frac{1}{z_2 + \Pi} \frac{N_1}{N_2} + \frac{1}{z_2 + \Pi} \frac{d\Pi}{dT_1} - \frac{(1-\kappa)\phi - \frac{\alpha(1-\phi)}{1+\eta}}{1 - \frac{\alpha(1-\phi)}{1+\eta}} \frac{1}{N_2} \frac{dN_1}{dT_1} \end{aligned}$$

The sign of welfare change in each region obtains by plugging in (B.23) and (B.24):

$$\begin{aligned} \text{sign} \left\{ \frac{dv_1}{dT_1} \right\} &= \text{sign} \left\{ N_1 \frac{z_1 - z_2}{z_2 + \Pi} (1-\kappa)\phi + \sigma \right\} \\ \text{sign} \left\{ \frac{dv_2}{dT_1} \right\} &= \text{sign} \left\{ N_2 \frac{z_1 - z_2}{z_1 + \Pi} (1-\kappa)\phi - \sigma \right\} \end{aligned} \quad (\text{B.25})$$

Case 1: Agglomeration \leq Congestion In this case $(1-\kappa)\phi \leq \frac{\alpha(1-\phi)}{1+\eta}$. We can safely have $\sigma \rightarrow 0$ while retaining a unique stable equilibrium. When $\sigma \rightarrow 0$, welfare changes are strictly positive in both regions since $\lim_{\sigma \rightarrow 0} N_1 > 0$ and $\lim_{\sigma \rightarrow 0} N_2 > 0$ and Π is bounded from above by Π^{\max} . It is Pareto-improving to introduce a small subsidy $dT_1 > 0$ and the small tax $dT_2 < 0$ that balances the budget. When $\sigma > 0$, derivative in region 1 is always positive so that the only Pareto improving candidate policy is $dT_1 > 0$. One can find a sufficient condition for the laissez faire to be efficient, i.e., a condition under which derivative in 2 is always strictly negative:

$$\sigma > (1-\kappa)\phi \frac{z_1 - z_2}{z_1}$$

To see this, use (B.25) and the fact that $N_2 < 1$ and $\Pi > 0$.

Case 2: Agglomeration > Congestion In this case $(1 - \kappa)\phi > \frac{\alpha(1-\phi)}{1+\eta}$. One can only allow a minimum of $\sigma \rightarrow (1 - \kappa)\phi - \frac{\alpha(1-\phi)}{1+\eta} > 0$ to retain a unique equilibrium. In this limiting case we have

$$\begin{aligned} \text{sign} \left\{ \lim_{\sigma \rightarrow 0} \frac{dv_1}{dT_1} \right\} &= \text{sign} \left\{ (1 - \kappa)\phi - \frac{\alpha(1 - \phi)}{1 + \eta} + (1 - \kappa)\phi \lim_{\sigma \rightarrow 0} N_1 \frac{z_1 - z_2}{z_2 + \Pi} \right\} > 0 \\ \text{sign} \left\{ \lim_{\sigma \rightarrow 0} \frac{dv_2}{dT_1} \right\} &= \text{sign} \left\{ \lim_{\sigma \rightarrow 0} \left(N_2 \frac{z_1 - z_2}{z_1 + \Pi} - 1 \right) (1 - \kappa)\phi + \frac{\alpha(1 - \phi)}{1 + \eta} \right\} \end{aligned} \quad (\text{B.26})$$

A sufficient condition for the impossibility of any Pareto improving reform when $\sigma \rightarrow (1 - \kappa)\phi - \frac{\alpha(1-\phi)}{1+\eta}$ — hence whatever the value of $\sigma > (1 - \kappa)\phi - \frac{\alpha(1-\phi)}{1+\eta}$ — is that agglomeration forces are too strong relative to congestion forces:

$$(1 - \kappa)\phi > \frac{z_1}{z_2} \frac{\alpha(1 - \phi)}{1 + \eta}$$

To see this, use (B.25) and the fact that $N_2 < 1$ and $\Pi > 0$. Note that with CRS technology for supplying the local non-traded good ($\eta \rightarrow \infty$, i.e., perfectly elastic housing supply), no Pareto improving reforms exists.

B.3 Planner’s Problem

Given exogenous Pareto weights $\{\omega_i\}_i$, the optimal allocation is the vector

$$\{v_j, c_j, h_j, G_j, N_j, \mathcal{O}_j, L_j^Y, L_j^G, L_j^H, L_j^{\mathcal{O}}\}_j$$

that maximizes

$$W = \mathbb{E}[\omega_i \ln u_i(v_1, \dots, v_J)]$$

Note that the planner is constrained to give the same c_j , h_j and G_j to all individuals of j hence the same v_j . The Lagrangian for the planner’s problem is

$$\begin{aligned}
L &= \mathbb{E} [\omega_i \ln u_i(v_1, \dots, v_J)] \\
&- \sum_j \lambda_j N_j [v_j - v_j(c_j, h_j, \mathbf{G}_j(G_j, N_j), \mathcal{G}(\{\mathcal{O}_k\}_k))] \\
&- \sum_j E_j \left[N_j - \frac{v_j^{1/\sigma}}{\sum_k v_k^{1/\sigma}} \right] \\
&- \pi^Y \sum_j [N_j c_j - F_j^Y(L_j^Y)] \\
&- \sum_j \pi_j^G [G_j - F_j^G(L_j^G)] \\
&- \sum_j \pi_j^{\mathcal{O}} [\mathcal{O}_j - F_j^{\mathcal{O}}(L_j^{\mathcal{O}})] \\
&- \sum_j \pi_j^H [N_j h_j - F_j^H(L_j^H)] \\
&- \sum_j \pi_j^L [L_j^Y + L_j^G + L_j^H + L_j^{\mathcal{O}} - N_j]
\end{aligned} \tag{B.27}$$

The first term is the planner’s objective that incorporates individuals’ incentive constraints when recognizing that $\ln u_i = \max\{\ln v_j + \ln \mu_{ij}\}_j$. The second constraint is simply the definition of v_j . The third term is the aggregate incentive constraint that limits the planner’s freedom in assigning workers to cities. Note that $N_j \leq \frac{v_j^{1/\sigma}}{\sum_k v_k^{1/\sigma}}$ for all j guarantee that $\sum_j N_j \leq 1$. The remaining terms are resource and market clearing constraints. $\{\lambda_j, E_j, \pi^Y, \pi_j^H, \pi_j^G, \pi_j^L, \pi_j^{\mathcal{O}}\}_j$ are multipliers. Below we give the first-order conditions for $\{v_j, c_j, h_j, G_j, N_j, \mathcal{O}_j, L_j^Y, L_j^G, L_j^H, L_j^{\mathcal{O}}\}_j$.

First-order condition for v_j :

$$\sigma(\Omega_j - \lambda_j v_j) + E_j - \sum_k N_k E_k = 0 \tag{B.28}$$

where

$$\Omega_j \equiv \frac{1}{N_j} \mathbb{E} [\omega_i \mathbf{1}_{[v_{ij} = \max\{v_{ik}\}_k]}] \tag{B.29}$$

First-order conditions for c_j, h_j and G_j :

$$\begin{aligned}\lambda_j \frac{\partial v_j}{\partial c_j} &= \pi^Y \\ \lambda_j \frac{\partial v_j}{\partial h_j} &= \pi_j^H \\ \lambda_j N_j \frac{\partial v_j}{\partial G_j} \frac{\partial G_j}{\partial G_j} &= \pi_j^G\end{aligned}\tag{B.30}$$

First-order conditions for L_j^Y, L_j^G, L_j^H and L_j^O :

$$\pi^Y \frac{\partial F_j^Y}{\partial L_j^Y} = \pi_j^G \frac{\partial F_j^G}{\partial L_j^G} = \pi_j^H \frac{\partial F_j^H}{\partial L_j^H} = \pi_j^O \frac{\partial F_j^O}{\partial L_j^O} = \pi_j^L\tag{B.31}$$

First-order condition for N_j :

$$\lambda_j N_j \frac{\partial v_j}{\partial G_j} \frac{\partial G_j}{\partial N_j} - E_j - (\pi^Y c_j + \pi_j^H h_j) + \pi_j^L = 0\tag{B.32}$$

First-order condition for O_j :

$$\sum_k \lambda_k N_k \frac{\partial v_k}{\partial G} \frac{\partial G}{\partial O_j} = \pi_j^O\tag{B.33}$$

To get to expression (2.32) in the text, plug (B.30) and (B.31) in expression (B.32). To get to expression (2.33), plug (B.30) and (B.31) in expression (B.33).

B.4 Public Good Demand Calibration

In order to get a sense of the value of the congestion parameter κ , we follow the seminal approaches of [Borchering and Deacon \(1972\)](#) and [Bergstrom and Goodman \(1973\)](#) and calibrate a model of local public good demand. In Section 2.4, we model how non-myopic jurisdictions maximize residents’ utility $v_j(c_j, h_j, G_j)$ over head tax τ_j^h and property tax τ_j^p knowing the per capita government transfer T_j subject to the municipality budget constraint, and embed this optimization in a general equilibrium framework with mobile households. We reproduce here the local Samuelson rule that emerges:

$$\frac{p_j^G}{N_j} \frac{\partial v_j}{\partial c_j} = p \frac{\partial v_j}{\partial G_j} \frac{\partial G_j}{\partial G_j}\tag{B.34}$$

To bring equation (B.34) to empirical analysis, we assume a generalized constant elasticity

of substitution shape for v_j :

$$v_j(c_j, h_j, \mathbf{G}_j) = (\alpha_C c_j^{-\rho_C} + \alpha_H h_j^{-\rho_H} + \alpha_G \mathbf{G}_j^{-\rho_G})^{1/\rho} f_j(X_j)$$

where c_j and h_j are consumption of traded and local non-traded goods, $\mathbf{G} = \frac{G}{N^\kappa} \frac{1}{h(T)}$ is congested local public goods, $f_j(X_j)$ is a taste shifter and

$$\rho = \frac{1 - \sigma}{\sigma}$$

$$\sigma = \frac{\theta_C}{1 + \rho_C} + \frac{\theta_H}{1 + \rho_H} + \frac{\theta_G}{1 + \rho_G}$$

with $\{\theta_C, \theta_H, \theta_G\}$ being consumption, housing and public spending budget shares. Generalized CES utility approximates constant non-unitary price and income elasticities provided that budget shares do not vary too much (see [Sato 1972](#)). Under this constant share assumption, the Samuelson rule ([B.34](#)) directly translates into

$$\rho_G \alpha_G (\mathbf{G}_j)^{-\rho_G - 1} = \rho_C \alpha_C \frac{p_j^G}{p} \frac{1}{N_j^{1-\kappa}} (c_j)^{-\rho_C - 1} f_j(X_j)$$

Using the constant share assumption for c_j and multiplying by p_j^G , we get the following public good expenditure function:

$$\frac{E_j}{N_j} = \frac{p_j^G G_j}{N_j} \equiv (p)^{\frac{1}{1+\rho_G}} (p_j^G)^{\frac{\rho_G}{1+\rho_G}} (N_j)^{\frac{\rho_G}{1+\rho_G}(\kappa-1)} (w_j + T_j + \Pi)^{\frac{1+\rho_G}{1+\rho_G}} T_j^\gamma f_j(X_j) \quad (\text{B.35})$$

which allows to separately calibrate ρ_G and κ , knowing the value of ρ_C . Note that $E_j = p_j^G G_j$ where $p^G = \left(\frac{\zeta p_s^G}{1-\psi}\right)^{1-\psi} \left(\frac{p_f^G}{\psi}\right)^\psi$ when $G = G_s^\psi G_f^{1-\psi}$ and where ζ is a constant capturing the fact that durable public goods are financed over time through debt repayment. Many variables should be observed by the econometrician. In particular, equation ([B.35](#)) justifies using a weighted geometric average of observed public spending on the left-hand side even though real outputs are unobserved.

We bring expression ([B.35](#)) to regression analysis after taking logs. As we observe all 2,000 municipal federations representing 36,000 French municipalities in all even years between 2002 and 2014, we use the panel structure of our data to identify the relationship between per capita spending and within-federation changes in population and other confounders. Time-invariant local confounders over 2002-2014 are absorbed by MF fixed effects in f_j , while time-specific remaining shocks in local prices are absorbed by flexible region-specific time effects. To account

for residual time variation within MF, we control for local housing prices and local wages. Finally, we flexibly control for within-MF time variation in industry sector and occupational composition X_j .¹

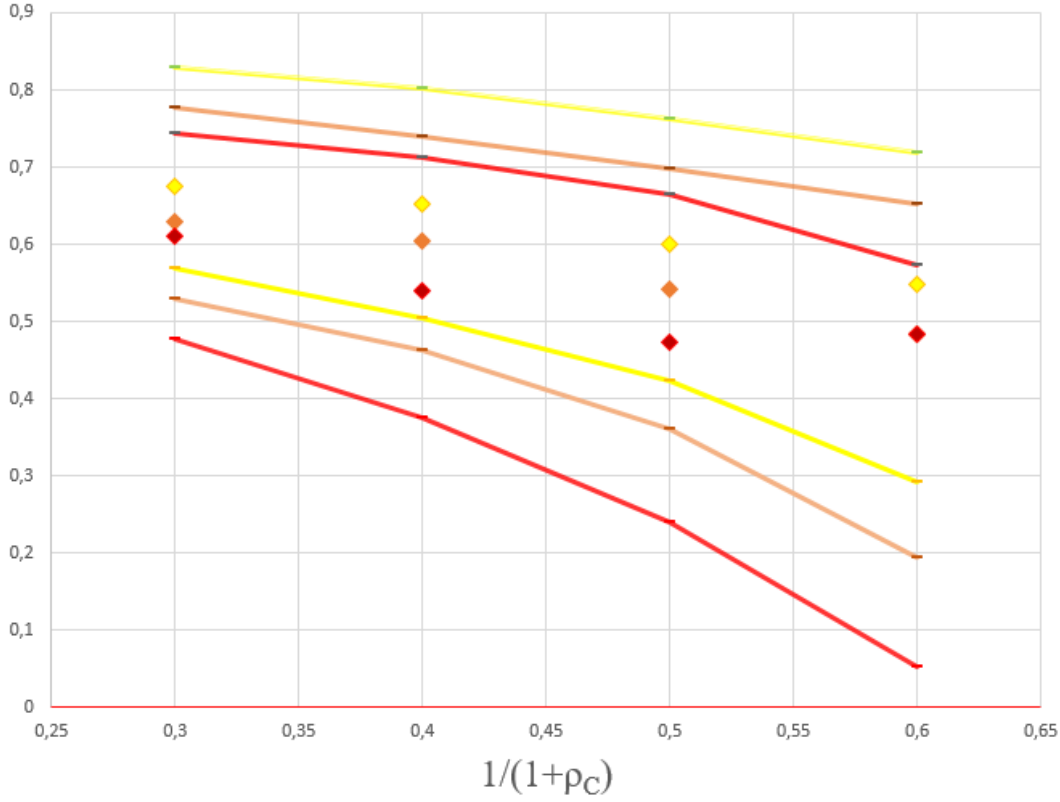
To derive “structura” elasticities from the reduced form estimates of equation (B.35), we use estimates for consumption price elasticity $\frac{1}{1+\rho_C}$ from studies on French households. We compute an average price-elasticity of consumption without housing of $\frac{1}{1+\rho_C} = 0.32$ from Faure et al. (2012) and $\frac{1}{1+\rho_C} = 0.62$ from Abramovici (1994). We report structural elasticities for $\frac{1}{1+\rho_C} = 0.5$ in Table B.1 and carry out a sensibility analysis of parameter κ for values of $\frac{1}{1+\rho_C} \in \{0.3, 0.4, 0.5, 0.6\}$ in Figure B.1. Our calibrated κ is between 0.44 and 0.59, meaning that estimates for public good agglomeration elasticity $1 - \kappa$ lie between 0.41 and 0.56.

Table B.1: Structural Elasticities for $\frac{1}{1+\rho_C} = 0.5$

	(1)	(2)	(3)
Price Elasticity $\left(\frac{1}{1+\rho_G}\right)$	0.478	0.487	0.521
	[0.278;0.685]	[0.271;0.708]	[0.273;0.745]
Congestion Elasticity (κ)	0.586	0.517	0.444
	[0.395;0.747]	[0.305;0.682]	[0.224;0.629]
FE	✓	✓	✓
Controls: Local Prices	✓	✓	✓
Controls: Industry & Occupation		✓	✓
Year FE \times			MA

Note: This Table reports estimates of structural elasticities. We flexibly control for the share of the population in various industry sectors and occupations. MA stands for metropolitan area. Bias-corrected point estimates and 95% confidence intervals are computed with a wild cluster bootstrap procedure based on 1,000 resamplings of residuals, where clusters are MAs.

¹We predicted MF outcomes net of individual characteristics. However, industry sector and occupational composition bears a geographic component which is independent of individuals and may still affect local demand for public goods.

Figure B.1: Sensibility Analysis for κ 

Note: This graph shows the sensibility of our estimation of κ to both the hypothesis on ρ_C and the chosen specification in Table B.1 : Yellow stands for (1), Orange for (2) and Red for (3), with ρ_C varying from 0.3 to 0.6.

B.5 Constant Expenditure Shares

The use of a generalized CES utility in the model for public good demand exposed and appendix B.4 relies on the constancy of expenditure shares approximation. Using housing and public good expenditures present in our dataset, we take this hypothesis to an empirical test. Exact rent r_j^h dataset being unavailable outside of the main French metropolitan areas, we use here discounted housing prices per square meter r_j as a measure of rent. Public expenditures are defined as in Section 2.3 and scaled by the residents contribution share to municipal budget τ .

$$E_j^h = h_j \cdot r_j = h_j \frac{R}{1+R} p_j^h$$

$$E_j^g = \tau \left(\zeta p_{s,j}^G G_{s,j} + p_f^G G_{f,j} \right)$$

Table B.2: Cross-Sectional Estimates

	(1)	(2)
	ln(Housing Expenditure p.c.)	ln(Public Good Expenditures p.c.)
log(Income p.c.)	1.036*** (0.264)	0.960*** (0.246)
Controls	✓	✓
Year FE X	UA	UA
p-value $H_0 : (\beta = 1)$	0.891	0.871
R-squared	0.789	0.616
Observations	7,731	7,768

Note: UA stands for metropolitan area. Standard errors are clustered at the UA level. * significant at 10%; ** significant at 5%; *** significant at 1%.

Table B.3: Linear Regression Share Estimates

	(1)	(2)
	Housing Expenditure p.c.	Public Good Expenditures p.c.
Income p.c.	0.137*** (0.028)	0.035*** (0.009)
Controls	✓	✓
Year FE X	UA	UA
R-squared	0.790	0.610
Observations	7,731	7,768

Note: UA stands for metropolitan area. Controls are industry & occupation. Standard errors are clustered at the UA level. * significant at 10%; ** significant at 5%; *** significant at 1%.

Expenditure shares can be considered constant throughout cities for a representative individual if the hypothesis of linearity of income in per capita expenditure across cities, which writes $H_0 : \beta = 1$, cannot be rejected for the following cross-sectional specification:

$$\ln E_{jt} = \alpha + \beta \ln y_{jt} + X_{jt} \cdot \gamma + \lambda_{at} + \epsilon_{jt}$$

in which we allow for metropolitan area specific time trends and control for city socio-demographic characteristics X_{jt} . Tables B.2 reports the results of the test, from which it appears that the constant share hypothesis is consistent with our data. Table B.3 reports the results of a linear specification which coefficients are the shares of expenditures in each sector. Estimates of housing expenditure share are smaller than existing estimates (typically around 30%), but recall that we include subsidies to local governments in our definition of total disposable income.

B.6 Calibration of the Public Good Index

In this section, we calibrate the parameter of the public good index ψ . Table B.4 and Figure B.2 test the relevance of our Cobb-Douglas specification. Table B.5 estimates the Cobb-Douglas parameter by solving $\frac{1-\psi}{\psi\zeta} = \frac{1}{0.212}$ where $\zeta = 0.049$ accounts for public asset depreciation and the fact that public assets are financed through debt repaid over time as in Jannin and Sotura (2019). We estimate $\hat{\psi} = 0.81$.

Table B.4: Log-Linearity Test

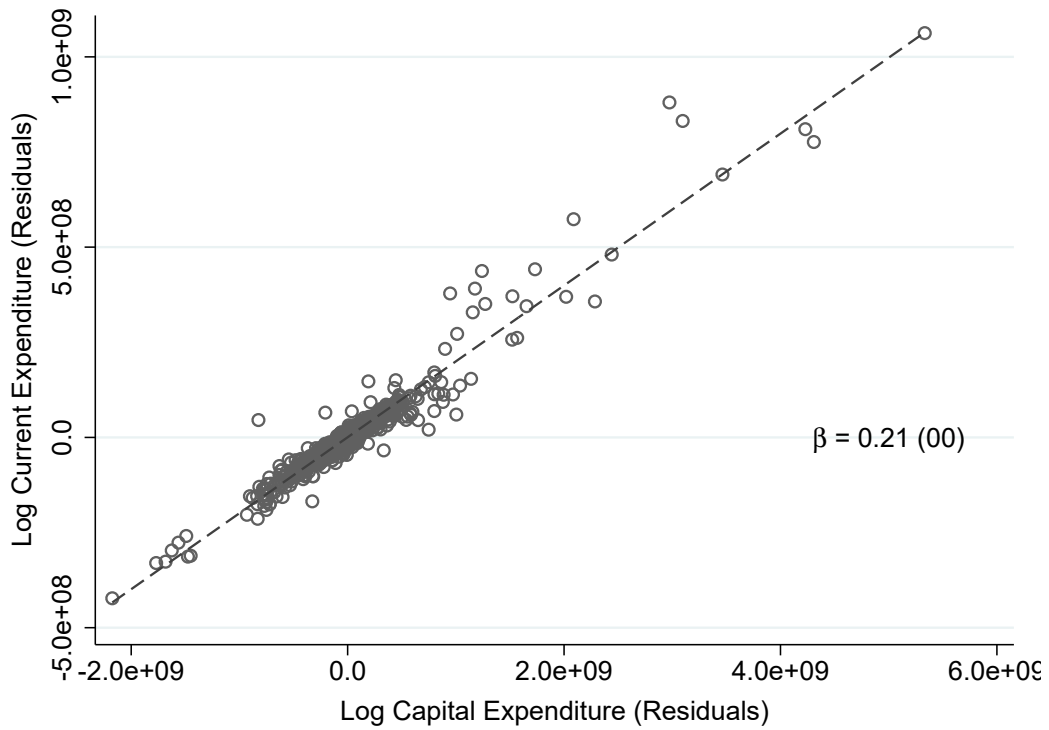
	log(Operating Expenditure)
log(Capital Expenditure)	1.024*** (0.009)
Controls	✓
Year FE X	UA
R-squared	0.966
Observations	9,710

Note: UA stands for metropolitan area. Standard errors are clustered at the UA level. * significant at 10%; ** significant at 5%; *** significant at 1%.

Table B.5: Calibration of ψ

	Operating Expenditure
Capital Expenditure	0.212*** (0.010)
Controls	✓
Year FE X	UA
R-squared	0.957
Observations	9,710

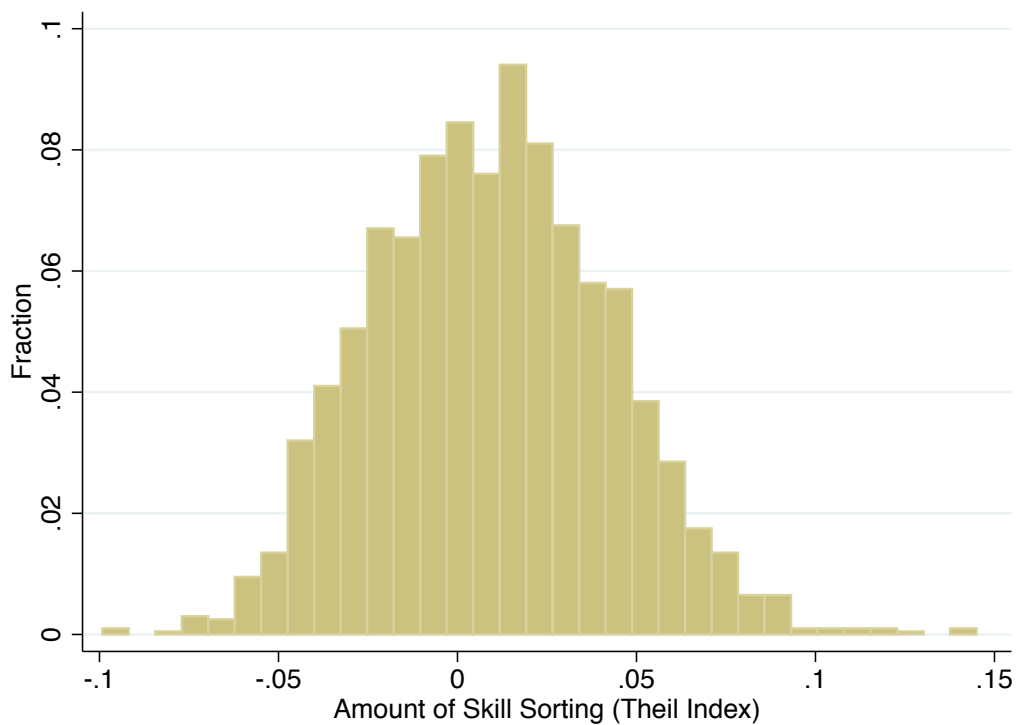
Note: UA stands for metropolitan area. Standard errors are clustered at the UA level. * significant at 10%; ** significant at 5%; *** significant at 1%.

Figure B.2: Calibration of ψ 

Note: This graph shows the relationship between log current expenditure and log capital expenditure, both residualized with respect to $UA \times year$ fixed effects. Current expenditure are the sum of yearly staff expenditure, maintenance spending, payments for external services and operating subsidies to third parties. Capital expenditure are the current book value of durable facilities and are the sum of all public assets such as schools, transportation infrastructure, parks improvements, sports facilities, museums, art collections, investment subsidies to local clubs, minus the raw value of the land and financial assets.

B.7 Skill Sorting

Figure B.3: Skill Sorting



Note: This graph shows the amount of skill sorting across MFs in 2008. The Theil is defined for each MF as $\sum_k \text{share}_k \times \ln(\text{share}_k / \text{share}_k^{\text{ref}})$ where share_k is the 2008 share of the MF population with education level $k \in \{1, 2, 3, 4\}$ and $\text{share}_k^{\text{ref}}$ the corresponding share in national population. Perfect sorting would yield MFs with Theil around $-\ln(\text{share}_k^{\text{ref}}) \in [2.2; 2.8]$ depending on education level k . No sorting yields MFs with Theil around 0. As can be seen, skill sorting is very low in our setting.

Appendix C

Appendix to “The Deadweight Loss of Property Transaction Taxes”

C.1 Comparative Statics

Equilibrium is characterized by tightness and expected price $\{\theta^*(\tau), p^{e*}(\tau)\}$ such that

$$\begin{aligned}\theta &= \frac{\bar{v}(\theta, p^e)}{\bar{e}(\theta, \tau)} \\ p^e &= \mathbb{E}[p|\theta, \tau] = \int_{\mathcal{B}(\tau)} \frac{e_i(\theta, \tau)}{\bar{e}(\theta, \tau)} p_i(\tau)\end{aligned}\tag{C.1}$$

Totally differentiate the first line of (C.1):

$$d\theta = \frac{1}{\bar{e}} (\bar{v}_1 d\theta + \bar{v}_2 dp^e) - \bar{v} \frac{1}{\bar{e}^2} (\bar{e}_1 d\theta + \bar{e}_2 d\tau)\tag{C.2}$$

where subscripts denote derivative with respect to arguments. We can express partial derivatives:

$$\begin{aligned}\frac{\partial \theta}{\partial p^e} &= \frac{\bar{v}_2}{\bar{e}} \left(1 - \frac{\bar{v}_1}{\bar{e}} + \frac{\bar{v}}{\bar{e}^2} \bar{e}_1\right)^{-1} > 0 \\ \frac{\partial \theta}{\partial \tau} &= -\bar{e}_2 \frac{\bar{v}}{\bar{e}^2} \left(1 - \frac{\bar{v}_1}{\bar{e}} + \frac{\bar{v}}{\bar{e}^2} \bar{e}_1\right)^{-1} > 0\end{aligned}\tag{C.3}$$

We also have

$$\begin{aligned}\frac{\partial p^e}{\partial \theta} &= \int_{\mathcal{B}(\tau)} \frac{\partial (e_i/\bar{e})}{\partial \theta} p_i \geq 0 \\ \frac{\partial p^e}{\partial \tau} &= \int_{\mathcal{B}(\tau)} \underbrace{\frac{e_i}{\bar{e}} \frac{\partial p_i}{\partial \tau}}_{<0} + \int_{\mathcal{B}(\tau)} \underbrace{\frac{\partial (e_i/\bar{e})}{\partial \tau} p_i}_{\geq 0} \geq 0\end{aligned}\tag{C.4}$$

C.2 Proof of Proposition 1

Totally differentiate p^e with respect to τ using $p^e(\theta, \tau)$ and $\theta(p^e, \tau)$:

$$\frac{dp^e}{d\tau} = \frac{\partial p^e}{\partial \tau} + \frac{\partial p^e}{\partial \theta} \left(\frac{\partial \theta}{\partial \tau} + \frac{\partial \theta}{\partial p^e} \frac{dp^e}{d\tau} \right)$$

so that

$$\frac{dp^e}{d\tau} = \left(\frac{\partial p^e}{\partial \tau} + \frac{\partial p^e}{\partial \theta} \frac{\partial \theta}{\partial \tau} \right) \left(1 - \frac{\partial \theta}{\partial p^e} \frac{\partial p^e}{\partial \theta} \right)^{-1} \quad (\text{C.5})$$

With homogeneous agents $e_i = \bar{e}$ so that $\frac{\partial p^e}{\partial \theta} = 0$ and

$$\frac{dp^e}{d\tau} = \frac{\partial p^e}{\partial \tau} = \int_{\mathcal{B}(\tau)} \frac{e_i}{\bar{e}} \frac{\partial p_i}{\partial \tau} < 0$$

hence prices always decrease with the tax.

With heterogeneous agents and under Assumption 1 (low types are more responsive):

$$\frac{dp^e}{d\tau} = \left(\underbrace{\frac{\partial p^e}{\partial \tau}}_{\geq 0} + \underbrace{\frac{\partial p^e}{\partial \theta} \frac{\partial \theta}{\partial \tau}}_{\leq 0} \right) \underbrace{\left(1 - \frac{\partial \theta}{\partial p^e} \frac{\partial p^e}{\partial \theta} \right)^{-1}}_{> 0} \geq 0 \quad (\text{C.6})$$

so that prices may increase or decrease with the tax depending on the strength of the composition channel, i.e., the second term of

$$\frac{\partial p^e}{\partial \tau} = \int_{\mathcal{B}(\tau)} \frac{e_i}{\bar{e}} \underbrace{\frac{\partial p_i}{\partial \tau}}_{< 0} + \int_{\mathcal{B}(\tau)} \underbrace{\frac{\partial (e_i/\bar{e})}{\partial \tau}}_{\geq 0} p_i \geq 0$$

C.3 Proof of Proposition 2

With search costs $c_b(e) = e^{\gamma_b}$ and $c_s(e) = e^{\gamma_s}$ ($\gamma_{b,s} > 1$):

$$\begin{aligned} e_i &= [\alpha_b(\theta)\beta (b_i - \tau - s)]^{\frac{1}{\gamma_b-1}} \\ v &= [\alpha_s(\theta) (p^e - s)]^{\frac{1}{\gamma_s-1}} \end{aligned} \quad (\text{C.7})$$

With matching technology $n(\bar{e}, \bar{v}) = \bar{e}^\eta \bar{v}^{1-\eta}$ one shows that

$$\theta(p^e, \tau) = (p^e - s)^{\frac{\zeta}{\gamma_s-1}} \beta^{\frac{\zeta}{1-\gamma_b}} \times \left(\int_{\mathcal{B}(\tau)} (b_i - \tau - s)^{\frac{1}{\gamma_b-1}} \right)^{-\zeta} \quad (\text{C.8})$$

with $\zeta = \left(1 + \frac{\eta}{\gamma_s - 1} + \frac{1 - \eta}{\gamma_b - 1}\right)^{-1} > 0$. It is immediate to see that $\frac{\partial \theta}{\partial p^e} > 0$ since only the first term depends on p^e . Furthermore, $\theta(s, \tau) = 0$ and $\lim_{p^e \rightarrow +\infty} \theta(p^e, \tau) = +\infty$. Call $\tilde{p}^e(\theta, \tau)$ the inverse function of $\theta(p^e, \tau)$, strictly increasing with θ with $\tilde{p}^e(0, \tau) = s$ and $\lim_{\theta \rightarrow +\infty} \tilde{p}^e(\theta, \tau) = +\infty$.

Notice that $\frac{d}{db_i} \frac{\partial \ln e_i}{\partial \theta} = 0$ so that $\frac{\partial p^e}{\partial \theta} = 0$ and $p^e(\theta, \tau) > s$ does not depend on θ . Hence by continuity $p^e(\theta, \tau)$ and $\tilde{p}^e(\theta, \tau)$ cross only once and above s at the equilibrium tightness $\theta^*(\tau)$. Equilibrium price $p^{e*}(\tau) = p^e(\theta, \tau)$ is fully characterized by the tax level. Note that $p^{e*}(\tau)$ encompasses both bargaining and composition responses.

C.4 Proof of Proposition 3

Differentiate equation (3.22) and plug first order conditions for buyers' and sellers' conditional search effort:

$$\frac{dW}{d\tau} = \frac{dn}{d\tau} (\mathbb{E}[b|\tau] - s) + n \frac{d\mathbb{E}[b|\tau]}{d\tau} - n (\mathbb{E}[p|\tau] - s) \frac{1}{\bar{v}} \frac{d\bar{v}}{d\tau} - n \int_{\mathcal{B}(\tau)} \beta [b_i - \tau - s] \frac{1}{\bar{e}} \frac{de_i}{d\tau} \quad (\text{C.9})$$

where I used $n = \alpha_b \bar{e} = \alpha_s \bar{v}$. I neglect changes in $\mathcal{B}(\tau)$ since marginal buyers exert zero effort and $c_b(0) = 0$. Plugging in the price expression and reorganizing yields

$$\begin{aligned} \frac{dW}{d\tau} &= \frac{dn}{d\tau} (\mathbb{E}[b|\tau] - s) + n \frac{d\mathbb{E}[b|\tau]}{d\tau} - n (\mathbb{E}[b|\tau] - \tau - s) (1 - \beta) \frac{1}{\bar{v}} \frac{d\bar{v}}{d\tau} \\ &\quad + n(\tau + s) \beta \frac{1}{\bar{e}} \frac{d\bar{e}}{d\tau} - n \beta \int_{\mathcal{B}(\tau)} \frac{1}{\bar{e}} \frac{de_i}{d\tau} b_i \end{aligned}$$

Note that

$$d \int_{\mathcal{B}(\tau)} \frac{e_i}{\bar{e}} b_i = \int_{\mathcal{B}(\tau)} \frac{de_i}{\bar{e}} b_i - \frac{d\bar{e}}{\bar{e}} \int_{\mathcal{B}(\tau)} \frac{e_i}{\bar{e}} b_i$$

and equivalently that

$$\int_{\mathcal{B}(\tau)} \frac{de_i}{\bar{e}} b_i = d\mathbb{E}[b|\tau] + \frac{d\bar{e}}{\bar{e}} \mathbb{E}[b|\tau]$$

Hence

$$\begin{aligned} \frac{dW}{d\tau} &= \frac{dn}{d\tau} (\mathbb{E}[b|\tau] - s) + n \frac{d\mathbb{E}[b|\tau]}{d\tau} - n (\mathbb{E}[b|\tau] - \tau - s) (1 - \beta) \frac{1}{\bar{v}} \frac{d\bar{v}}{d\tau} \\ &\quad + n(\tau + s) \beta \frac{1}{\bar{e}} \frac{d\bar{e}}{d\tau} - n \beta \frac{d\mathbb{E}[b|\tau]}{d\tau} - n \beta \frac{1}{\bar{e}} \frac{d\bar{e}}{d\tau} \mathbb{E}[b|\tau] \\ &= \frac{dn}{d\tau} (\mathbb{E}[b|\tau] - s) + (1 - \beta) n \frac{d\mathbb{E}[b|\tau]}{d\tau} \\ &\quad - n (\mathbb{E}[b|\tau] - \tau - s) \left[\beta \frac{1}{\bar{e}} \frac{d\bar{e}}{d\tau} + (1 - \beta) \frac{1}{\bar{v}} \frac{d\bar{v}}{d\tau} \right] \end{aligned}$$

Using

$$\frac{dn}{d\tau} = \frac{\partial n}{\partial \bar{e}} \frac{d\bar{e}}{d\tau} + \frac{\partial n}{\partial \bar{v}} \frac{d\bar{v}}{d\tau}$$

and Assumption 2 (Hosios) one gets

$$\frac{1}{n} \frac{dn}{d\tau} = \beta \frac{1}{\bar{e}} \frac{d\bar{e}}{d\tau} + (1 - \beta) \frac{1}{\bar{v}} \frac{d\bar{v}}{d\tau}$$

such that

$$\frac{dW}{d\tau} = \tau \frac{dn}{d\tau} + (1 - \beta)n \frac{d\mathbb{E}[b|\tau]}{d\tau}$$

which completes the proof.