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Gábor Kátay

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## THÈSE DE DOCTORAT

# Essais sur deux enjeux majeurs des pays d'Europe de l'Est: l'endettement en devises étrangères et l'offre de travail

Gábor Kátay

20 novembre 2015

Directeur de thèse: Fabrizio Coricelli

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# Résumé en français

La chute du rideau de fer a créé une occasion historique d'unifier l'Europe. Les anciens pays socialistes d'Europe de l'Est se sont engagés dans un processus de transition d'une économie planifiée centralisée à une économie de marché, ce qui impliquait des changements institutionnels et structurels fondamentaux. Alors que ces pays ont été confrontés à des défis similaires au cours des premières années de la transition, il y a eu d'importantes disparités dans l'approche politique et dans l'impact de ces politiques adoptées sur la performance économique de ces pays. Vingt-cinq ans plus tard, plusieurs Pays d'Europe Centrale et Orientale souffrent de certaines faiblesses communes, tandis que d'autres ont eu plus de succès à mettre leurs économies sur la voie d'une croissance durable. L'expérience de ces pays, qu'elle soit positive ou négative, fournit des enseignements précieux à la fois pour la communauté scientifique et pour les politiciens.

Cette thèse traite deux sujets distincts, représentant chacun des enjeux importants pour un grand nombre de pays d'Europe de l'Est, dont les origines remontent au changement de régime et aux réponses politiques successives. La première partie de la thèse porte sur les emprunts en devises étrangères. En utilisant une base de données détaillées au niveau du contrat du prêt bancaire pour les sociétés non-financières hongroises, le premier chapitre explore le choix des entreprises entre la monnaie locale et plusieurs devises étrangères dans une situation où les taux d'intérêt du marché d'une ou plusieurs devises étrangères sont plus bas que celle de la monnaie locale. La seconde partie est consacrée à l'exploration des liens entre les systèmes socio-fiscaux et l'offre de travail à la marge extensive. Le deuxième chapitre présente une nouvelle stratégie de modélisation de l'offre de travail à la marge extensive comme alternative aux deux approches dominantes basées sur le calcul marginal et sur les modèles d'utilité aléatoire. Finalement, le dernier chapitre utilise ce modèle pour quantifier la part de la différence entre les taux d'activité hongrois et tchèques qui peut être attribuée à des différences dans les deux systèmes fiscal-sociaux. Dans ce qui suit, je présente brièvement les deux questions dans une perspective historique et internationale, puis je résume les conclusions principales des papiers de recherche correspondants.



## L'endettement en devises étrangères

Depuis le début des années '90, le rattrapage économique dans les pays post-socialistes a été porté par des investissements directs importants en provenance des économies développées et par une forte expansion du crédit. La demande soutenue de crédit et les contraintes de liquidité du secteur bancaire dans ces pays a encouragé les banques de l'Europe occidentale à pénétrer les marchés des pays d'Europe de l'Est au moyen d'acquisitions de banques détenues par l'Etat pendant les programmes de privatisation ou des investissements "greenfield". L'investissement étranger dans le secteur bancaire a été mutuellement bénéfique pour l'investisseur et le marché de destination : les pays d'Europe de l'Est ont bénéficié d'un accès plus facile aux fonds externes, tandis que les investisseurs ont obtenu des taux de rendement élevés.

Compte tenu du faible niveau initial de l'intermédiation financière et d'une croissance potentielle relativement élevée, la hausse substantielle du ratio de crédit sur PIB était perçue comme un approfondissement financier soutenu favorisant la convergence économique de cette région. Bien que ce fût certainement le cas au cours des premières années de la transition, il est maintenant évident que la hausse du crédit surpassait le taux soutenable dans certains de ces pays. En outre, les risques de crédit ont été aggravés par la généralisation des emprunts libellés en devises étrangères, ce qui a exposé les ménages ainsi que les entreprises (sans couverture naturelle) aux risques de change.

En effet, dans de nombreux pays d'Europe de l'Est, l'emprunt en devises étrangères par les ménages et les entreprises a été la norme plutôt que l'exception. Avant l'éclatement de la crise, la majorité de la dette des entreprises a été libellée en devises étrangères en Bulgarie, Hongrie, Lituanie, Macédoine, Roumanie, Serbie et Slovaquie. Dans une moindre mesure, l'endettement en devises étrangères était également répandu dans d'autres pays d'Europe de l'Est. L'emprunt en devises étrangères par les ménages affiche une hétérogénéité plus importante entre les pays : il a été sollicité par une part croissante des ménages hongrois, lituaniens, roumains et serbes, tandis que l'emprunt en devises étrangères par les ménages a été pratiquement nul en Croatie, en République tchèque et en Slovaquie.

Les préoccupations relatives à la stabilité financière ont été évoquées par plusieurs banques centrales et institutions internationales bien avant la crise, mais elles ont été soit ignorées, soit dévalorisées. Même si elles étaient reconnues, de nombreux économistes et politiciens ont anticipé (ou espéré) un ajustement progressif vers une trajectoire soutenable. Avec l'éclatement de la crise financière, ces attentes ou espoirs se sont révélés largement illusoire. La crise a mis en évidence la vulnérabilité de plusieurs pays d'Europe de l'Est fortement endettés en devises étrangères, ce qui a contribué à la forte récession économique de la région.

En réponse au risque accru d'une crise systémique à l'échelle régionale et de contagion financière du système bancaire des pays avancés, l'Initiative de Vienne a été lancée en Janvier 2009 afin de réunir les institutions internationales compétentes, les autorités réglementaires et fiscales, ainsi que les plus grandes banques opérant dans la région. L'initiative a contribué à résoudre le classique dilemme du prisonnier en coordonnant une réponse commune. En effet, pour se protéger contre le risque de défaut croissant, les grandes banques étrangères retiraient leurs fonds de leurs filiales dans les pays touchés. Bien que cette stratégie non coopérative soit parfaitement rationnelle du point de vue individuel de chaque banque, des retraits de fonds à grande échelle menaçaient d'une crise financière systémique qui aurait affecté tout le monde. L'Initiative de Vienne a réussi à faire respecter l'engagement des grandes banques et des investisseurs à maintenir leur exposition dans ces pays et ainsi stabiliser le secteur financier de la région.

La motivation du premier chapitre de la thèse découle naturellement de l'importance de ce problème. Pour évaluer les risques associés à l'endettement excessif en devises étrangères, il est avant tout impératif de comprendre les décisions d'emprunt en devises étrangères des entreprises. Le premier chapitre étudie la volonté des entreprises d'apparier la composition en devises de leurs actifs et leurs passifs ainsi que leurs incitations à dévier de l'appariement parfait. Afin de motiver notre analyse empirique, nous présentons un modèle théorique simple et transparent et nous dérivons une solution en forme fermée pour le portefeuille optimal de la dette. En utilisant une base de données détaillées sur les contrats des sociétés non-financières hongroises, nous construisons – en cohérence avec la théorie – un indicateur de “non-appariement” et nous testons son influence sur le choix des entreprises de libeller leurs emprunts en devises étrangères. L'importance relative des motifs liés à la couverture naturelle versus d'autres facteurs tels que l'écart des taux d'intérêt est également abordée. Finalement, nous explorons les effets de la crise sur le choix de devise des entreprises non-financières.

Pour tester l'existence du motif d'appariement dans le choix de devise des entreprises, les études empiriques existantes relient généralement la part de la dette en devises étrangères, d'une part, à un proxy de la sensibilité des profits des entreprises aux fluctuations de taux de change et, d'autre part, à d'autres variables individuelles ou macroéconomiques. Les premières études examinent la décision de la dénomination de la dette des grandes entreprises aux Etats-Unis (Allayannis and Ofek (2001), Kedia and Mozumdar (2003)) et en Finlande (Keloharju and Niskanen (2001)). Ces études trouvent que les entreprises dont la part des exportations (ou la fraction des actifs détenus à l'étranger) est plus élevée que la moyenne détiennent relativement plus de dette libellée en devises étrangères afin de couvrir leur exposition aux risques de change. Un nombre considérable d'études sur les marchés émergents – dont beaucoup sont publiées dans un numéro spécial de *Emerging Market Review* (vol. 4 no. 4) – tendent vers la même conclusion. Dans ces études, l'identification repose sur la part des exportations ou sur un indicateur de biens commercialisables (variable in-

diquant si l'entreprise appartient au secteur marchand) pour montrer que les exportateurs ou les entreprises produisant des biens échangeables sont davantage susceptibles de détenir des dettes en devises étrangères. Le coefficient de ces variables étant généralement positif et – à l'exception des résultats argentins et brésiliens – significativement différent de zéro, les auteurs concluent que les entreprises ont tendance à appairer la composition en devises de leurs dettes avec la sensibilité de leur chiffre d'affaires au taux de change.

Dans le même temps, le rôle du différentiel de taux d'intérêt et donc le motif de spéculation ("carry trade") est une raison souvent citée pour la dollarisation dans les études macro. Rosenberg and Tirpak (2009) examinent les déterminants du recours croissant à l'euro et au franc suisse dans les nouveaux Etats membres de l'Union Européenne. Les auteurs trouvent que le différentiel de taux d'intérêt est une variable explicative robuste de différences dans les taux d'endettement en devises étrangères entre les pays. Basso et al. (2011) montrent que l'écart de taux d'intérêt est un facteur important pour la dollarisation des prêts et des dépôts.

Le premier chapitre de la thèse met en évidence trois lacunes importantes de la littérature empirique existante. Premièrement, les preuves existantes montrant que les entreprises exportatrices ont tendance à emprunter davantage en devises étrangères ne fournissent pas de preuve directe du motif d'appariement. L'exportation est seulement la moitié de l'histoire : la mesure appropriée pour tester les considérations d'appariement entre les actifs et les passifs doit inclure non seulement les revenus d'exportation libellés en devises étrangères mais aussi les paiements de la dette en devises étrangères. Par exemple, une entreprise qui s'endette en devise étrangère à un point tel que ses revenus d'exportation attendus ne couvrent pas entièrement ses obligations de paiement de la dette est exposée à un risque de change similaire à une entreprise non-exportatrice avec des passifs libellés en devise étrangère dans son bilan. De même, une entreprise dont la part des exportations est relativement faible peut s'endetter entièrement en devise étrangère en toute sécurité tant que ses revenus d'exportation restent supérieurs ou égaux à ses obligations de paiement de la dette.

Deuxièmement, l'hétérogénéité non observée des entreprises est rarement prise en compte. L'identification des motifs d'appariement par la comparaison de deux groupes distincts d'entreprises, les exportateurs et les non-exportateurs, qui sont probablement différents à beaucoup d'égards ne fournit pas de preuve directe pour l'appariement. Si le choix de libeller la dette en devise étrangère ou en monnaie locale est le résultat d'une optimisation, la même entreprise dans différentes circonstances doit faire des choix différents. Autrement dit, l'identification des motifs d'appariement n'est possible qu'une fois les effets fixes d'entreprise pris en compte.

Troisièmement, il n'y a pas actuellement de consensus clair quant à l'effet dominant dans la décision des entreprises. Seules quelques études se concentrent à la fois sur la couverture naturelle et les stratégies de carry trade. Parmi les quelques exceptions figurent les contributions de Keloharju

and Niskanen (2001) et de Brown et al. (2010), qui confirment que les deux motifs sont présents en même temps. En examinant les entreprises finlandaises et bulgares, respectivement, les auteurs constatent que les entreprises exportatrices sont davantage susceptibles de solliciter des prêts libellés en devises étrangères. En même temps, les entreprises ont également tendance à choisir la devise étrangère lorsque le différentiel de taux d'intérêt entre la devise étrangère et la monnaie locale est plus élevé qu'en moyenne. Une étude descriptive par Endresz et al. (2012) suggère également que les motifs d'appariement et de carry trade sont des facteurs pertinents : la dette libellée en devises étrangères est surtout concentrée parmi les plus grandes entreprises exportatrices, cependant, beaucoup de petites entreprises non-exportatrices se sont également endettées en devises étrangères en Hongrie. L'importance relative de ces deux facteurs est seulement abordée dans Brown et al. (2011). Les auteurs constatent que l'emprunt en devises étrangères est beaucoup plus fortement lié aux revenus d'exportation des entreprises qu'aux différentiels de taux d'intérêt, ce qui les pousse à conclure que la spéculation n'est pas le facteur clé de l'endettement en devises étrangères des entreprises.

L'expérience récente de la Hongrie en matière d'endettement en devises étrangères et une nouvelle collection de données administratives appariées – qui comprend les comptes et les bilans de toutes les entreprises hongroises, les revenus mensuels d'exportation et les dépenses d'importation, des informations sur tous les contrats de prêts aux entreprises et sur le fournisseur de crédit – offrent une opportunité unique de réévaluer les déterminants des décisions d'emprunt des entreprises.

Notre stratégie d'identification repose sur l'intuition que si l'appariement est un facteur pertinent, la probabilité de solliciter un nouveau prêt libellé en devise étrangère doit diminuer dès que l'obligation de remboursement de la dette en devises étrangères sur une période donnée dépasse les revenus d'exportation en devises étrangères attendues par l'entreprise. Cela conduit à un modèle à variable dépendante binaire dans lequel la probabilité de choisir la devise étrangère est reliée à un indicateur de non-appariement prenant la valeur 1 si les obligations de remboursement de la dette en devises étrangères sont plus élevées que les revenus d'exportation de l'entreprise et 0 dans le cas contraire. Dans un deuxième temps, une mesure plus précise de l'indicateur de non-appariement conduit à une spécification logit mixte (Train (2003)) dans laquelle des équations séparées pour les deux devises étrangères majeures – correspondant essentiellement à l'euro et au Francs suisses – sont estimées simultanément. Nous utilisons ensuite le modèle estimé pour effectuer une analyse contrefactuelle. Afin d'isoler l'effet du motif de couverture naturelle dans le choix de devise des entreprises, nous "éliminons" cet effet en mettant l'indicateur de non-appariement à 1 pour toutes les observations et ainsi nous prédisons les choix hypothétiques de devise des entreprises en l'absence de motifs de couverture.

Les résultats d'estimation sur un échantillon avant la crise montrent que la probabilité de contracter une dette en devise étrangère diminue dès que l'obligation de remboursement de la dette de l'entreprise en devises étrangères dépasse ses revenus d'exportation attendus. Ce résultat est robuste sous diverses spécifications du modèle et sous divers choix de l'échantillon, ce qui fournit une preuve solide à l'appui du rôle de la couverture naturelle dans le choix de la monnaie dans laquelle la dette des entreprises est libellée. Le motif de la couverture naturelle est encore plus fort depuis le déclenchement de la crise financière actuelle. Néanmoins, nos résultats suggèrent que ce dernier n'est pas le motif principal d'endettement en devise étrangère: il n'explique qu'environ 10 pour cent de la dette totale en devises étrangères des entreprises avant la crise, et 20 pour cent pendant la crise. La plus grande partie de la dette en devises étrangères correspondrait, au moins en Hongrie, à des positions de carry trade détenues par des sociétés non financières.

En outre, nos résultats indiquent que les entreprises avec une forte probabilité de choisir un endettement en euro ont également une forte probabilité de choisir un endettement en franc suisse. Les avantages de détenir de la dette à la fois en euro et en franc suisse semblent donc l'emporter sur l'avantage de choisir systématiquement une devise étrangère (celle préférée par l'entreprise) à l'autre, ce qui peut être interprété comme un indice du fait que les entreprises accordent plus d'importance à la diversification qu'à la pleine exploitation des opportunités d'arbitrage (carry trade) entre les devises étrangères.

L'étude montre également que depuis le déclenchement de la crise actuelle, l'attractivité relative du franc suisse s'est considérablement détériorée non seulement vis-à-vis de la monnaie locale, mais aussi vis-à-vis de l'euro. En conséquence, un certain nombre de sociétés ont abandonné le franc suisse en partie pour la monnaie locale, mais également en partie pour l'euro. Ce changement dans les préférences relatives entre les monnaies étrangères suggère que la stratégie de carry trade entre les deux devises étrangères aurait joué un rôle dans les décisions stratégiques des entreprises.

## **Systèmes socio-fiscaux et l'offre de travail**

L'impact du changement de régime sur le marché du travail a été profond dans tous les pays d'Europe de l'Est. Le résultat presque inévitable de la transition depuis des emplois garantis vers un marché du travail gouverné par l'offre et la demande est la forte hausse du chômage. D'une part le chômage déguisé en emploi s'est transformé en chômage réel. D'autre part, la restructuration de l'économie, exigeant de nouvelles compétences, a causé un problème sérieux d'appariement sur le marché du travail. Alors que certaines professions ont connu de graves pénuries de main-d'œuvre, la plupart des personnes âgées et moins éduquées n'ont pas pu acquérir les compétences nécessaires pour répondre à la nouvelle demande.

Après le changement de régime, la plupart des pays d'Europe de l'Est se sont rapidement engagés dans des réformes dans de nombreux domaines tels que la gouvernance, la privatisation, ainsi que la libéralisation des prix, du commerce et des marchés de capitaux. Cependant, les réformes sur le marché du travail étaient plus difficiles et plus lentes à mettre en œuvre. Les gouvernements étaient pris entre deux objectifs contradictoires : d'une part, la dérégulation du marché du travail par la facilitation des licenciements et par la réduction des prestations universelles généreuses de l'ère socialiste ont suscité l'espoir d'une stabilisation macroéconomique et d'une convergence des revenus accélérées ; d'autre part, dans de nombreux pays, les préoccupations concernant les coûts sociaux de la transition et les engagements des gouvernements à soutenir les personnes touchées empêchaient la mise en œuvre des réformes à grande échelle sur le marché du travail.

Les poids relatifs des deux facteurs et les politiques qui en découlaient variaient considérablement d'un pays à l'autre. Les pays baltes étaient assez agressifs dans la déréglementation du marché du travail. En revanche, les pays d'Europe du Sud ont maintenu des politiques de protection de l'emploi relativement strictes. Cette politique prudente a contribué à réduire les tensions sur le marché du travail, mais au prix d'une création d'emplois moindre dans le secteur privé à productivité élevée. En Hongrie, au moins une partie de la réduction d'emploi a été compensée par l'encouragement de la retraite anticipée, la facilitation de l'accès au régime de pension d'invalidité et par d'autres prestations sociales. Une politique similaire a été adoptée en Pologne, ce qui a contribué à la baisse du taux de chômage élevé aux dépens du taux d'activité.

La performance du marché du travail d'un certain nombre de pays de la région est toujours à la traîne derrière l'UE15. Alors qu'en moyenne, les taux d'activité et le taux d'emploi sont généralement plus faibles dans les nouveaux états membres que dans les pays d'Europe occidentale, les pays qui ont entrepris des réformes audacieuses du marché du travail pendant les premières années de la transition, tels que l'Estonie et la Lettonie, semblent être récompensés sur le long terme. A l'autre extrême, la Croatie et la Hongrie souffrent toujours d'un faible taux d'activité, même si les performances de leurs marchés du travail se sont progressivement améliorées ces derniers temps.

Dans l'ensemble, les différences des taux d'activité de ces pays et les différences dans les systèmes socio-fiscaux adoptés suggèrent que les choix des politiques fiscales et sociales ont des répercussions importantes sur la performance du marché du travail. Soutenues par une littérature abondante, les mauvaises incitations financières au travail sont généralement considérées comme les causes principales du faible taux d'activité par rapport à une norme internationale ou d'autres pays similaires. En conséquence, les organisations internationales telles que le FMI, l'OCDE ou la Commission Européenne recommandent souvent aux pays à faible taux d'activité de mettre en œuvre des réformes fiscal-sociales en vue d'améliorer les incitations au travail.

La deuxième partie de la thèse a deux objectifs : (1) proposer une nouvelle stratégie de modélisation de l'offre de travail comme alternative aux deux approches dominantes basées sur le calcul marginal et les modèles d'utilité aléatoire ; (2) en utilisant ce modèle, quantifier la part de la différence entre les taux d'activité tchèque et hongrois qui peut être expliquée par les divergences des systèmes d'imposition et de protection sociale.

Dans le modèle de base linéaire de Hausman, l'individu choisit un ensemble de consommation et d'heures travaillées qui maximise son utilité sous la contrainte budgétaire. Le résultat de ce problème de maximisation se traduit par une fonction d'offre de travail. Dans ce cadre, l'offre de travail dépend du salaire net, des revenus non-salariaux et des préférences individuelles. Les impôts affectent négativement l'offre de travail par l'intermédiaire de l'effet de substitution (les impôts diminuent le salaire net, ce qui diminue le coût d'opportunité de loisirs) et positivement par l'intermédiaire de l'effet de revenu (une baisse de salaire net a tendance à inciter les individus à travailler plus afin de compenser, au moins partiellement, la perte de revenu). La participation au travail rémunéré découle d'une simple solution extrême.

Le modèle de première génération de Hausman a été inlassablement critiqué pour plusieurs raisons. Tout d'abord, le modèle traditionnel de Hausman repose sur l'hypothèse de linéarité (par morceaux) et de convexité de l'ensemble de budget. Cette hypothèse est clairement violée si les individus perdent certaines prestations dès qu'ils commencent à travailler, ce qui est généralement le cas. En conséquence, le salaire gagné au cours des premières heures travaillées ne compense pas la perte des prestations et, par conséquent, le salaire de réserve est indéfini. Cette limitation est également valable pour des techniques d'estimation non paramétriques (Blomquist and Newey (2002)). Le modèle peut tenir compte des coûts fixes liés à la recherche d'emploi, mais cette source supplémentaire de non-convexité pose un nouveau défi qui est difficile à résoudre (voir par exemple Bourguignon and Magnac (1990)). van Soest and Das (2001) affirment que les non-convexités impliquent généralement des formes de préférences trop restrictives et peu plausibles.

Deuxièmement, le modèle de Hausman suppose implicitement la quasi-concavité de la fonction d'utilité, ce qui est généralement rejetée par les études empiriques (voir par exemple Blundell and MaCurdy (1999)). Comme indiqué par MaCurdy et al. (1990), le modèle de Hausman impose implicitement des restrictions qui génèrent un effet de Slutsky positif sur l'ensemble des points internes de la contrainte budgétaire. Par conséquent, la cohérence du modèle limite implicitement la gamme des élasticités qui peut être obtenue.

Troisièmement, les estimations issues du modèle de Hausman présentent une hétérogénéité notable et donnent souvent des mesures d'élasticité peu fiables tant à la marge extensive qu'à la marge intensive. Pour les hommes mariés, beaucoup d'études trouvent une élasticité compensée (hicksienne) négative des heures travaillées au salaire réel (voir Pencavel (1987) pour une revue des

résultats pour les hommes). En revanche, l'élasticité non compensée de l'offre de travail féminin est, dans de nombreux cas, irréaliste, souvent largement supérieure à un (voir Killingsworth and Heckman (1987) pour une revue des résultats pour les femmes) .<sup>1</sup> Bien que la comparaison directe des résultats entre les études soit souvent difficile en raison de différences méthodologiques – par exemple, la façon dont les salaires net sont traités diffèrent d'une étude à l'autre – et de différences dans le choix des sous-groupes de population traités, la méta-analyse de Chetty et al. (2013) fournit un "consensus" sur l'élasticité hicksienne de l'offre de travail à la marge extensive de 0,25. Cependant, à la marge intensive, le modèle de Hausman s'avère incapable de prédire avec précision la distribution des heures travaillées observée car il ne tient pas compte du fait que le nombre d'individus travaillant peu d'heures est très limité.

Finalement, ces études se concentrent habituellement sur soit les impôts, soit les transferts sociaux. Comme Blundell (2012) le souligne, il est important de prendre en compte simultanément les impôts et les transferts sociaux et de les combiner en des charges fiscales effectives. En plus d'influencer les revenus non salariaux, les transferts sociaux montrent également des caractéristiques de taux d'imposition à la fois marginal et moyen. Par exemple, considérons une prestation conditionnelle aux revenus avec une élimination progressive. Chaque euro supplémentaire de revenu réduit le montant de cette prestation de 20 centimes. Ceci est équivalent à un taux marginal d'imposition supplémentaire de 20%. Lorsque l'individu perd entièrement ses droits à cette prestation, cette perte devient semblable à un taux moyen d'imposition : le montant total des transferts sociaux perdus diminue la récompense du travail, tout comme le taux moyen d'imposition.

Depuis la publication des travaux fondateurs de van Soest (1995), l'approche de choix discret pour modéliser l'offre de travail est devenue de plus en plus populaire. Le modèle de choix discret est basé sur le concept de maximisation de l'utilité aléatoire dans lequel les individus choisissent une seule option parmi un ensemble fini d'heures possibles. Le modèle inclut l'inactivité (zéro heure travaillée) comme l'une des options. Ainsi, l'offre de travail à la marge extensive et intensive sont estimés simultanément. Contrairement à l'approche de Hausman, cette méthode ne modélise pas explicitement l'offre de travail. La fonction d'utilité est directement estimée et l'offre de travail des individus et des ménages est dérivée de cette fonction. En principe, le modèle est compatible avec un très large éventail de spécifications de préférences. De plus, les conditions de tangence et la convexité des préférences ne doivent pas être imposées. Le modèle ne nécessite que l'imposition de la monotonie de la consommation et la spécification précise des préférences pour la consommation et le loisir. En outre, il est relativement facile de prendre en compte la présence simultanée des impôts

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<sup>1</sup>Voir également la revue de littérature plus récente de Evers et al. (2008). Les articles de Blundell and MaCurdy (1999), Blundell et al. (2007), Keane (2011) et de Meghir and Phillips (2011) renvoient également la littérature empirique de l'approche continue de Hausman, mais incluent également des résultats plus récents sur la méthodologie des choix discrets.



sur les revenus et des transferts sociaux. Néanmoins, comme Dagsvik et al. (2014) le soulignent, le modèle de choix discret est similaire à l'approche standard d'un point de vue théorique car les deux sont issus de la même théorie du comportement du consommateur. La seule nouveauté est que l'ensemble des heures possibles est finie.<sup>2</sup>

Notre modèle contraste avec à la fois l'approche hausmanienne et celle du choix discret. En adoptant une théorie très similaire à l'approche de choix discret standard, nous dérivons l'expression de "gains de travail" d'acceptation, ce qui se traduit par une fonction d'offre de travail. La spécification est similaire à la fonction d'offre de travail issue du calcul marginal de modèle hausmanien, mais elle permet de tenir compte des non-linéarités de l'ensemble de choix. Dans un sens, nous fournissons un pont entre l'approche standard basée sur le calcul marginal et celle de la nouvelle approche de choix discret.

Dans notre modèle, la catégorie de référence est nécessairement l'inactivité et l'utilité dérivée de toute autre alternative est comparée à l'utilité de rester en dehors du marché du travail. Dans un premier temps, nous limitons le choix des heures travaillées à zéro et l'emploi à temps plein. Dans ce cas, l'individu choisit de travailler si ses gains de travail, définies comme la différence entre le salaire net et le montant des transferts sociaux perdus lorsque l'individu travaille, sont plus élevés que ses "gains de travail d'acceptation". La théorie sous-jacente conduit à une équation à variable dépendante binaire qui relie la probabilité de participation aux gains de travail, au montant total des revenus non salariaux (y compris le montant des transferts sociaux que l'individu obtient ou obtiendrait à zéro heure travaillée) et à d'autres caractéristiques individuelles. Les gains de travail et les revenus non-salariaux sont influencés par les impôts et les transferts sociaux conditionnels aux revenus. De même que dans le cas du modèle de choix discret standard, l'expression analytique pour les probabilités de sélection en présence de plusieurs choix possibles d'heures travaillées est dérivée en utilisant l'axiome d'indépendance des alternatives non pertinentes (IIA) (voir McFadden (1974)). Introduit par Luce (1959), l'IIA dit que les probabilités relatives entre deux choix possibles sont indépendantes de la présence ou l'absence d'autres alternatives. Sous cette hypothèse, les décisions de l'offre de travail à la marge extensive peuvent être représentées par un modèle logistique multinomial. L'hypothèse forte de l'IIA peut être assouplie en spécifiant un modèle de logit mixte, ce qui permet que les facteurs non observés soient corrélés entre les choix alternatifs.

Pour tester notre modèle, nous effectuons nos estimations sur l'enquête sur le budget des ménages hongrois (HKF). Notre base de données contient des mesures détaillées sur les revenus et la consommation des individus pour les années 1998-2008. La Hongrie présente un cas particulièrement

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<sup>2</sup>En parallèle des modèles structurels de l'offre de travail, il existe également une vaste littérature sur la méthodologie d'évaluation de programme. Bien que de telles approches sont capables d'estimer précisément l'impact d'une réforme fiscal-social-particulière, elles ne sont pas appropriées pour évaluer l'impact de futurs scénarios (hypothétiques). Voir Moffitt (2002) pour une revue de ce volet de la littérature.

intéressant pour estimer les déterminants de l'offre de travail. Le taux d'activité est parmi les plus faibles dans l'Union Européenne: avec un chiffre de 61,5% en 2008, la Hongrie se classait avant dernière dans l'Union Européenne après Malte, 9,2 points de pourcentage inférieur à la moyenne des 28 pays de l'Union Européenne. Pour 2014, cet écart s'est réduit à 6,8 points de pourcentage. La faible participation au marché de travail a souvent été identifiée comme un des principaux obstacles à la convergence réelle (voir par exemple Kátay (2009)). Cependant, le travail à temps partiel en Hongrie est relativement rare : selon Eurostat, la part des salariés à temps partiel était inférieure à 5% pendant la période considérée. Par conséquent, nous supposons qu'il n'y a que deux états du marché du travail, actifs et inactifs, et nous testons uniquement le modèle avec ces deux choix possibles. Les recherches futures seront nécessaires pour tester le modèle avec plusieurs alternatives.

La méthode d'estimation suit la procédure en trois étapes souvent utilisée, comme par exemple dans Kimmel and Kniesner (1998). L'élément clé de l'identification est le choix des instruments, c'est-à-dire les variables qui n'ont pas d'impact (ou qui ont un impact négligeable) sur l'offre de travail directement, mais impactent les salaires et ainsi influencent la participation au marché du travail indirectement. Nous soutenons que les régions et l'expérience professionnelle sont de telles variables.

Nous trouvons qu'une seule équation peut expliquer une grande hétérogénéité dans les réponses des individus aux impôts et transferts sociaux : il existe des différences marquées dans les effets marginaux conditionnels entre les différents sous-groupes. Les sous-groupes les plus sensibles sont les peu qualifiés, les femmes (mariées) en âge de procréer et les individus plus âgés, tandis que les personnes diplômées sont pratiquement insensibles aux changements fiscaux et sociaux à la marge extensive. Nos estimations impliquent une élasticité de l'offre de travail globale de 0,28, ce qui est en ligne avec le "consensus" de 0,25 proposé par Chetty et al. (2013).

Le dernier chapitre examine – pour la première fois dans la littérature – dans quelle mesure les différences dans les taux d'activité entre les pays peuvent être expliquées par les divergences des systèmes d'imposition et de protection sociale. La littérature la plus proche utilise des données individuelles ou désagrégées transnationales soit pour étudier les effets d'un changement de politique socio-fiscale sur l'offre de travail, soit pour comparer l'évolution des taux d'activité entre les pays. Sans souci d'exhaustivité, certaines études empiriques récentes transnationales analysent les effets des réformes sociales hypothétiques (par exemple Immervoll et al. (2007)) ; d'autres étudient les effets des réformes fiscales (voir par exemple Colombino et al. (2000) ou Paulus and Peichl (2008)) ; des modèles de micro-simulation sont également utilisés pour estimer les préférences de redistribution des gouvernements (Blundell et al. (2009); Bargain et al. (2011)). D'autres études utilisent des données individuelles ou désagrégées pour décrire l'évolution de l'offre de travail

dans certains pays, soit pour toute la population (Balleer et al. (2009) ou Blundell et al. (2011)), soit pour une sous-population sélectionnée (par exemple Cipollone et al. (2013) pour les femmes). Cependant, aucune de ces études ou d'autres études empiriques existantes n'ont jusqu'à présent tenté d'expliquer les différences de taux d'activité entre les pays par les différences des systèmes socio-fiscaux.

Nous prenons l'exemple de deux pays, la République tchèque et la Hongrie. Ces deux pays fournissent un cas intéressant pour la comparaison. Ils présentent beaucoup de similitudes : leurs économies sont géographiquement proches, similaires en taille et en niveau de développement, et partagent en partie une histoire commune. En particulier, les deux économies ont connu le plein emploi jusqu'au changement de régime à la fin des années 1980. Malgré tous ces facteurs communs, les taux d'activité sont très différents dans ces deux pays. Alors que le taux d'activité de la Hongrie est parmi les plus bas de l'Union Européenne, le taux d'activité tchèque est proche de la moyenne de l'UE15.

Dans un premier temps, nous reproduisons sur des données tchèques l'estimation d'offre de travail réalisée sur données hongroises présentée dans le chapitre précédent. Nous utilisons ensuite ces deux estimations parfaitement comparables pour simuler comment le taux d'activité d'un pays changerait si ce pays adoptait le système d'imposition et de protection sociale de l'autre pays.

Nos résultats d'estimation donnent des élasticités d'offre de travail similaires, ce qui suggère que les préférences individuelles sont essentiellement identiques dans les deux pays. Ce résultat est également valable pour les sous-populations en fonction du niveau d'éducation, le sexe et l'état matrimonial. Les simulations montrent qu'environ la moitié de la différence dans les taux d'activité de la population âgée de 15 à 74 ans s'explique par les différences des systèmes socio-fiscaux. Les résultats sont quasi-symétriques, ce qui signifie que si le système tchèque était adopté, le taux d'activité de la Hongrie augmenterait d'environ le même nombre de points de pourcentage que le taux d'activité tchèque diminuerait si le système hongrois était mis en place. La catégorie qui verrait son taux d'activité le plus évoluer avec l'adoption du système fiscal-social de l'autre pays serait les femmes mariées ou en âge de procréer, les prestations de maternité étant beaucoup plus généreuses en Hongrie.

# General introduction

The fall of the Iron Curtain created a historic opportunity to unify Europe. The former socialist countries in Europe have been engaged in a transition process from centrally planned to market-oriented economy, involving fundamental institutional and structural changes. While these countries were facing similar challenges during the first few years of the transition, there have been significant disparities in the region in the policy approach and the impact of the policies adopted on the economic performance of the respective countries. Twenty-five years later, several Central and Eastern European (CEE) countries suffer from some common weaknesses, while others were more successful in setting their economies on a sustainable growth path. The experience of these countries, be it positive or negative, provides valuable lessons for both the scientific community and policy makers.

This thesis deals with two distinct topics, both of them representing central issues for many CEE countries which can be traced back to the regime change and the subsequent policy responses. The first part of the thesis focuses on (mainly corporate) foreign currency lending. Using detailed information at the loan contract level for the Hungarian non-financial corporate sector, the first chapter explores corporate borrowers' choice between the local currency and several possible foreign currencies in a situation where the market interest rates in one or several foreign currencies are lower than that of the local currency. The second part is devoted to exploring the links between tax-benefit systems and labour supply at the extensive margin (i.e., the labour force participation). The second chapter presents an alternative modelling strategy of labour supply to the two dominating approaches based on marginal calculus and on random utility models. Finally, the last chapter uses this model to quantify the difference between the Hungarian and the Czech participation rates that can be attributed to differences in taxation and welfare benefits. In what follows, I briefly present the two issues from a historical and international perspective, then I summarise the main findings of the corresponding papers.<sup>3</sup>

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<sup>3</sup>Chapter 1 presents a joint study with Péter Harasztosi. Chapter 2 is closely related to a previous paper joint with Péter Benczúr, Áron Kiss and Olivér Rácz. Finally, Chapter 3 represent the body of work I have done jointly with Kamil Galuščák. This latter research was supported by Czech National Bank Research Project No. D1/12.

## Foreign currency indebtedness

Most of the post-socialist countries made significant progress in convergence during the nineties and the beginning of this century. The catch-up process was further fueled by both significant direct capital flows from advanced economies to emerging markets and strong credit expansion. The buoyant demand for credit and the liquidity-constrained banking sector in these countries encouraged Western European banks to enter CEE markets by means of acquisitions of state-owned banks during the privatization programs or greenfield investments. The foreign investment in the banking sector was mutually beneficial for both the investor and the destination market: while CEE countries benefited from the easier access to external funds, investors achieved high rates of return.

Given the low initial level of financial intermediation and the relatively high potential for further growth, the substantial increase in the credit-to-GDP ratio was mainly perceived as sustained financial deepening which would foster economic convergence in the region. While this was certainly true during the first few years of the transition, it is now clear that credit growth rates surpassed the sustainable level in some of these countries. Moreover, the credit risks were compounded by the widespread use of foreign currency denominated borrowings, which exposed both households and firms (without natural hedge) to exchange rate shocks.

Indeed, lending in foreign currencies to both households and firms has been the norm in many CEE countries rather than the exception. As shown in the first chapter, before the outbreak of the crisis, the majority of the corporate debt was denominated in foreign currency in Bulgaria, Hungary, Lithuania, Macedonia, Romania, Serbia and Slovenia. To a lesser extent, the share of foreign currency in total debt was prevalent in other CEE countries as well. Foreign currency lending to the household sector displays a larger heterogeneity across countries. It has been mostly prevalent in Hungary, Lithuania, Romania and Serbia, while the share of FX borrowing by households has been virtually zero in Croatia, in the Czech Republic and in Slovakia.

Financial stability concerns have been raised by several central banks and international institutions already before the crisis, but they were either ignored or downplayed. Even if acknowledged, many economists and policy makers expected (or hoped for) a gradual adjustment towards a sustainable path. With the outbreak of the financial crisis, these expectations or hopes turned out to be largely illusory. The crisis brought to the surface the vulnerability of several CEE countries with high level of bank lending denominated in foreign currency, which contributed to the region's sharp economic downturn.

As a response to increasing risk of region-wide systemic crisis in the emerging European financial sector and the resulting financial contagion of the advanced countries' banking system, the so called "Vienna Initiative" had been launched in January 2009 in order to bring together the most impor-

tant European and international institutions, regulatory and fiscal authorities and the largest banks operating in the region. The initiative helped to solve a classic prisoner's dilemma by coordinating a common response. Indeed, to protect themselves against the increasing default risk, larger foreign banks had been pulling their funds out of their subsidiaries in the affected countries. Although this non-cooperative strategy is perfectly rational from the individual bank's perspective, large scale fund withdrawals threatened a full systemic financial crisis from which none would have escaped. The Vienna Initiative succeeded in upholding the commitment of the large banks and investors to maintaining exposure to subsidiaries and thereby stabilizing the financial sector in the region.

In light of the importance of the problem, the motivation for the first chapter of the thesis follows naturally. To evaluate the risks associated with excessive foreign currency indebtedness, it is crucial to understand firms' foreign borrowing decisions. The first chapter investigates firms' willingness to match the currency composition of their assets and their liabilities and their incentives to deviate from perfect matching. By adopting a simple mean-variance approach from modern portfolio theory, we first present a simple formal theoretical model on firms' optimal currency choice in the presence of multiple available foreign currency loans. In light with the proposed theory, we then rely on Hungarian corporate loan data and estimate discrete choice models in which firms choose the currency denomination of their loans.

Estimation results on a sample before the outbreak of the crisis show that the probability of borrowing in foreign currency decreases as soon as the firm's foreign currency debt reimbursement obligation exceeds its expected export revenues. This finding is robust across various model specifications and sample choices, which provides – for the first time in the literature – strong direct evidence to support the role of natural hedging motivation in firms' currency choice. Matching motivation is even stronger after the outbreak of the current financial crisis than during the pre-crisis period. However, our results suggest that natural hedging is not the main cause of firms' FX indebtedness: only about 10 percent of the overall corporate FX debt is attributable to natural hedging purpose during the pre-crisis and 20 percent during the post-crisis periods.

The motives for firms to deviate from the perfect matching portfolio is also addressed. In particular, we investigate corporate borrowers' choice between several possible foreign currencies – essentially euro and Swiss francs – and we test whether diversification strategies and/or carry-trade incentives between foreign currencies are relevant factors in firms' decisions. Results suggest that firms place higher value on diversification than on exploiting perceived arbitrage (or carry trade) opportunities, if any, between foreign currencies. At the same time, the paper shows that since the outbreak of the current crisis, the Swiss francs has lost its relative attractiveness vis-à-vis not only the local currency, but also vis-à-vis the euro. As a result, a number of firms have switched from the Swiss francs partly to the local currency, but also partly to the euro.

## Tax-benefit systems and labour supply

The labour market and thus the social impact of the regime change has been profound in all countries. The almost inevitable result of the transition from guaranteed employment to labour markets governed by supply and demand is the sharp increase in unemployment. Not only the earlier “in-door unemployed” lost their jobs. The restructuring of the economy required new skills, which caused a serious matching problem in the labour market. While there were serious shortages of suitable workforce in some professions, many of the elderly, less educated people could not acquire the necessary skills to meet the new demand.

Most of the CEE countries opted for front-loaded reforms (a.k.a. “shock therapy”) in many areas such as governance, privatization, liberalization of prices, trade and capital markets. Implementing labour market reforms, however, have been more difficult and protracted. Governments were caught between two conflicting objectives: first, the deregulation of the labour markets by easing layoffs and hiring procedures and the cut back of the generous universal benefits of the socialist era raised the hope for faster macroeconomic stabilization, recovery and income convergence; second, in many countries, concerns about the social costs of the transition and governments’ commitments to support affected individuals restrained the implementation of large-scale labour market reforms.

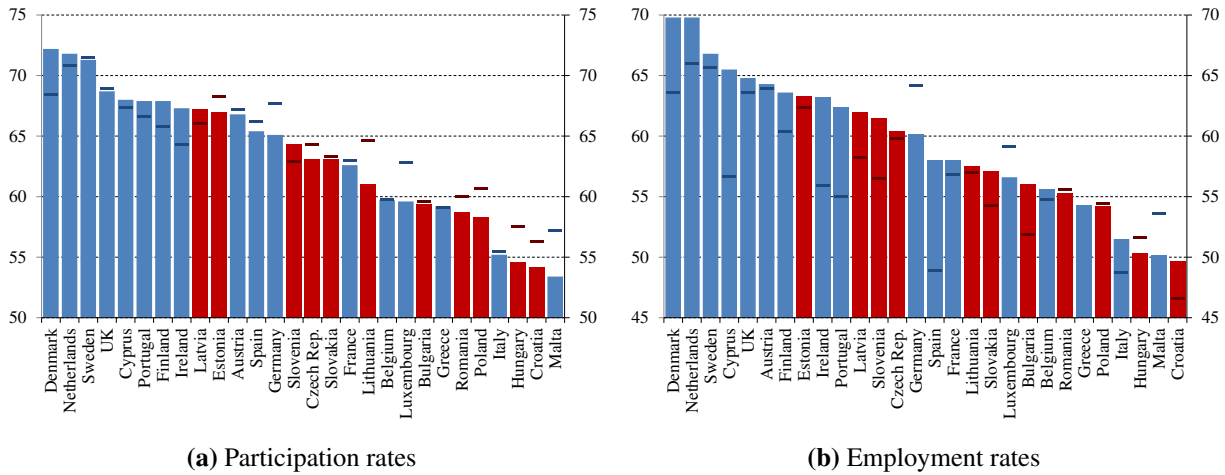
The relative weights of the two factors and the resulting policies varied markedly throughout the region. Baltic countries were fairly aggressive in early labour market deregulation. By contrast, Southeastern European countries maintained relatively strict employment protection policies. The cautious policy helped to reduce tensions on the labour market, but at a cost of less job creation in the higher productivity private sector.<sup>4</sup> In Hungary, at least part of the reduction in employment at the beginning of the nineties has been accommodated by encouraging early retirement, easing access to the disability retirement scheme and providing other welfare benefits. A similar policy has been adopted in Poland, which helped to reduce the high unemployment rate at the expense of declining labour force participation.

Labour market outcomes still lag behind the EU15 in a number of countries in the region. While on average, both participation rates and employment rates are generally lower in the New Member States than in Western Europe, countries that undertook bold labour market reforms at the early stage of the transition such as Latvia and Estonia seem to be rewarded in the longer run. At the other extreme, Croatia and Hungary still suffer from low participation and employment, even though their labour market outcomes have been gradually improved lately (see Figure 0.1).

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<sup>4</sup>At the same time, there is evidence of relatively weak enforcement of EPL in these countries.

**Figure 0.1.:** Participation and employment rates, 2008 & 2013



\* Notes: The figure shows participation rates (left figure) and employment rates (right figure) of the population aged 15-74 in 2008 (bars) and 2013 (lines). CEE countries are marked in red. Labour force survey data from Eurostat.

Overall, cross-country differences in labour participation and the differences in the adopted tax-benefit systems suggest that tax and welfare policy choices largely influence the country’s labour market outcome. Supported by a large body of literature, financial incentive to work is generally viewed as the primary candidate for explaining labour participation deficits of a given country relative to an international standard or against other similar countries. As a consequence, international organisations such as IMF, OECD or the European Commission frequently recommend for countries with low labour market participation to implement tax reforms and/or changes in the welfare system in order to improve work incentives.

The objective of the second part of the thesis is twofold. The first paper of this topic presents an alternative modelling strategy of labour supply to the existing approaches in the literature. Using individual (household level) data for Hungary and a multitude of tax and transfer reforms, we extend existing structural form methodologies and estimates the effect of income taxation and transfers on labour supply at the extensive margin. The second paper investigates the extent to which cross-country differences in aggregate participation rates can be explained by divergence in tax-benefit systems. We take the example of two countries, the Czech Republic and Hungary, which – despite a lot of similarities – differ markedly in labour force participation rates. For both countries, we use detailed microsimulation models and perfectly comparable micro estimates of labour supply to quantify the portion of the divergence in the two countries’ participation rates (in 2008) explained by differences in their taxation and welfare benefit systems.



In the early generation labour supply models relying on tangency conditions (Hausman (1981)), non-participation is simply determined by the difference between the market wage and the reservation wage. The framework is, however, restricted to the case of (piecewise) linear and convex budget sets. This assumption is particularly restrictive if certain transfers get lost immediately at taking up a job – which is usually the case – and the wage earned during the first few hours worked does not compensate for the discrete loss in benefits. On the other hand, the new “discrete choice approach” (van Soest (1995)) does not require tangency conditions to be imposed and the model is in principle very general. The model however does not explicitly estimate a labour supply function. Instead, utility functions are directly estimated and individuals’ labour supply is derived from individuals’ or households’ utility functions.

Our model contrasts with both the Hausman approach and the standard discrete choice modelling framework. Starting from a very similar theory as the standard discrete choice approach, we derive an expression for a “reservation gains to work”, which then translates into a labour supply function in which individuals’ labour supply choice depend on the financial gains to work, the non-labour income and other individual characteristics. We end up estimating a simple binary (or multinomial) choice specification with gains to work, non-labour income and individual characteristics being the covariates. The setup is similar to the standard textbook marginal calculus, but it is adopted to reflect the nonlinearities of the choice set. In a sense, we provide a bridge between the standard textbook approach based on marginal calculus and the new discrete choice approach of modelling labour supply.

Hungary provides a particularly interesting case for estimating the determinants of labour market activity. Participation rates are among the lowest in the EU (Figure 0.1). This has been often identified as a key bottleneck to real convergence. Moreover, our sample period featured numerous policy measures on both income tax rates and transfers, providing us with sufficient exogenous cross-sectional and time-series variation in non-labour income and gains to work.

We find that a single equation can explain a large heterogeneity of individual responsiveness to taxes and transfers: there are marked differences in the conditional marginal effects among subgroups. The most responsive subgroups are low-skilled, (married) women at child-bearing age and elders, while prime-age individuals with tertiary education are practically unresponsive to tax and transfer changes at the extensive margin.

The last chapter provides the first tentative explanation for cross-country differences in labour force participation rates attributed to differences in taxation and welfare benefits. We first replicate for Czech household-level data the labour supply estimation for Hungary and use the two perfectly comparable estimates to simulate how the aggregate participation rate would change in one country if the other country’s tax and social welfare system were adopted.

Our results show that the estimated labour supply elasticities for the Czech Republic are very close to the results for Hungary, suggesting that, at least in this dimension, individual preferences are similar in the two countries. This holds true even for sub-populations depending on the level of education, gender and marital status. Second, the simulation results suggest that about one-half of the total difference in the participation rates of the 15–74 years old population can be explained by differences in the tax-benefit systems. The results are quasi-symmetric, meaning that if the Czech system was adopted, the Hungarian participation rate would increase by about the same number of percentage points that the Czech participation would decrease by if the Hungarian system were implemented. The highest responses are obtained for married women and women of childbearing age, which is related to the more generous maternity benefit system in place in Hungary as compared to the Czech Republic.



# 1. Currency Matching and Carry Trade by Non-Financial Corporations

## 1.1. Introduction

Matching debt payments to expected foreign currency (FX) revenues is a natural way to mitigate the adverse effects of foreign exchange risk exposure of exporting firms. Currency mismatch occurs when firms' assets and liabilities are denominated in different currencies. Financial stability concerns typically arise when firms' net FX denominated liabilities are greater than their net FX denominated cash flows, i.e. when firms borrow "too much" in FX compared to their export revenues. As a result, if the domestic currency depreciates, firms with currency mismatch are likely to experience adverse balance sheet effects as the negative effect of the rise in FX denominated liabilities expressed in local currency usually outweighs the traditionally assumed positive competitiveness effect (Eichengreen et al. (2007)).

Currency mismatch is one of the main financial systemic risk factors that many emerging markets have been facing over the past decades. The debate about the importance of firms' currency mismatches and the adjustments following an unexpected exchange rate depreciation was particularly intense following the Latin American "Tequila" crisis. In these countries, the perceived security of the exchange rate peg encouraged liability dollarisation, while abandoning the peg resulted in a sharp depreciation of the respective home currencies. Due to the extensive dollarisation of firms' balance sheets and widespread currency mismatches, the weakening of the exchange rate and the implied increase in firms' debt burden in local currency terms forced firms to adjust their balance sheets, which translated into lower investment activities, declining production and, in many cases, liquidation proceedings. (see e.g. Krugman (1999) or Aghion et al. (2001)). Later, a similar process took place in Asia during the financial crisis of 1997-98 and in several Central and Eastern European countries following the outbreak of the current crisis.

This paper explores corporate borrowers' choice between the local currency and several possible foreign currencies in a situation where the market interest rates in one or several foreign currencies

are lower than that of the local currency. In particular, we investigate firms' willingness to match the currency composition of their assets and liabilities and their incentives to deviate from perfect matching. To motivate our empirical analysis, we use a simple and transparent model to derive a closed form solution for optimal debt portfolio. Using detailed information at the loan contract level for the Hungarian non-financial corporate sector, we then construct a theory-consistent measure of a mismatch indicator and we test its influence on firms' choice of currency denomination of borrowings. In doing so, this paper is the first to provide direct evidence to support the role of matching incentives in firms' debt currency choice. The relative importance of matching motivation versus other factors such as the interest rate differential is also addressed. Finally, we explore the effects of the crisis on firms' preferences and risk assessments.

To test for the existence of matching motivations in firms' debt currency choice, existing empirical studies usually relate the share of FX debt to a proxy for the sensitivity of firms' profits to exchange rate fluctuations and other firm specific or macroeconomic variables. The earliest studies investigate the currency-of-denomination decision of large US firms (Allayannis and Ofek (2001), Kedia and Mozumdar (2003)) and Finnish firms (Keloharju and Niskanen (2001)) and find that firms with higher export shares (or a larger share of total assets held abroad) hold more FX denominated debt in order to hedge their increased foreign exposure. A considerable number of studies on emerging markets – many of which are published in a special issue of the *Emerging Market Review* (vol. 4 no. 4) – point to the same conclusion.<sup>1</sup> These papers follow a similar reduced form estimation method in which the identification relies on the variable export share or on an indicator of tradability (dummy variable indicating whether the firm belongs to the tradable sector) to show that exporters or firms producing tradable goods are more likely to carry foreign debt. The coefficient of these variables being usually positive and – with the exception of the Argentinian and Brazilian results – significantly different from zero, the authors conclude that firms tend to match the currency composition of their liabilities with the ex-ante sensitivity of their revenues to the real exchange rate.

On the other hand, the role of the interest rate differential and thus carry trade behaviour is often put forward as an explanation for dollarisation in macro studies. Rosenberg and Tirpak (2009) examine the determinants of euroisation and swissfrancisation in the new EU Member States and find that the interest rate differential is a robust explanatory variable of cross-country differences. Basso

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<sup>1</sup>The special issue starts with a summary of Galindo et al. (2003) that collects results and findings from existing literature to that date. Existing studies on emerging markets cover Argentina (Galiani et al. (2003)), Brazil (Bonomo et al. (2003), Janot et al. (2008)), Chile (Benavente et al. (2003), Cowan et al. (2005), Fuentes (2009)), Colombia (Echeverry et al. (2003)), Mexico (Pratap et al. (2003), Gelos (2003), Martinez and Werner (2002)) Peru (Carranza et al. (2003)), Lebanon (Mora et al. (2013)) and several East Asian (Allayannis et al. (2003)) and Latin-American (Kamil (2012)) economies.

et al. (2011) find that the interest rate differential is a significant driver for the dollarisation of both loans and deposits.

This paper highlights and addresses three important shortcomings of the existing empirical literature. First, the existing evidence that exporting firms tend to borrow more in FX does not provide direct proof for currency matching. Export revenues are only half of the story, the appropriate measure for testing matching considerations between assets and liabilities should include both export revenues denominated in FX and FX debt payments. For instance, a firm that increases its FX debt to a point that the expected export revenues do not fully cover its debt payments is exposed to a similar exchange rate risk as non-exporters with FX liabilities in their balance sheets. Similarly, it is safe for a firm with a relatively low export share to incur debt entirely in FX as long as its export revenues are higher than or equal to its debt payment obligations.

Second, unobserved firm heterogeneity is rarely taken into account. Comparing two distinct groups of firms, (larger) exporters and (smaller) non-exporters, which are probably different in many respects, and identifying firms' matching incentives from cross-sectional variation does not provide direct evidence for matching. If the choice of currency denomination of the debt is a result of an optimisation process, the same firm in different circumstances should make different choices. That is, the identification of matching incentives is only possible once firm fixed effects are controlled for.

Third, there is as yet no clear understanding of which effect dominates in a firm's decision to incur FX-denominated debt. Only a few studies concentrate on both natural hedging and carry trade incentives. Among a few exceptions are the contributions by Keloharju and Niskanen (2001) and Brown et al. (2010), who confirm that both matching and carry trade motives are present in firms' currency choice. By examining Finnish and Bulgarian firms, the authors find that exporters are more likely to request FX loans, while firms also tend to choose FX when the interest rate differential between FX and domestic currency is higher than average. A descriptive study by Endresz et al. (2012) also suggests that both matching and carry trade are relevant factors: FX debt is mostly concentrated among larger, exporting firms, but many small non-trading firms also take out FX loans in Hungary. The relative importance of the two factors is only addressed in Brown et al. (2011). The authors find that FX borrowing is much stronger related to firm-level FX revenues than it is to country-level interest rate differentials, which makes them conclude that speculation is not the key driver of firms' currency choice.

The recent Hungarian experience of FX indebtedness together with a newly available collection of matched administrative datasets – which includes financial reports of all Hungarian firms, monthly export revenues and import expenditures, information on all corporate loan contracts and on the

credit provider – provide a valuable opportunity for reassessing the determinants of firms' borrowing decisions in a situation where they have access to FX loans.

Our identification strategy relies on the intuition that if currency matching is a relevant factor, the probability of incurring new FX-denominated debt must decrease as soon as the firm's FX debt reimbursement obligation during a given period of time exceeds its expected export revenues. This yields a binary dependent variable model in which the probability of choosing FX is related to a mismatch indicator which takes the value of 1 if the firm's debt payment obligations denominated in FX are higher than its export revenues and 0 otherwise. As a second step, a more accurate measure of mismatch indicator yields a mixed logit specification (Train (2003)) in which separate but correlated equations for two possible FX alternatives – corresponding essentially to euro (EUR) and Swiss francs (CHF) – are estimated simultaneously. We then use the estimated model to perform a counterfactual analysis. To isolate the effects of natural hedging motives on the aggregate corporate FX debt share, we “switch off” the effects of currency matching on firms' debt denomination choice by setting the mismatch indicator to 1 for all firms and for all choice occasions and we predict the counterfactual currency shares of newly contracted corporate loans in the absence of hedging motives.

The coefficient of the mismatch indicator being negative and highly significant across all specifications, our results provide strong evidence to support the role of natural hedging motivation in firms' currency choice. Matching motivation is even stronger in the aftermath of the crisis than during the pre-crisis period. However, our simulations indicate that natural hedging is not the primary motivation for firms to choose FX: it explains only about 10 percent of the overall corporate FX debt during the pre-crisis and 20 percent during the post-crisis periods.

In addition, our results indicate that firms with a higher probability of choosing EUR have also a higher probability of choosing CHF. The benefits of holding both EUR and CHF debt thus seems to outweigh the advantage of consistently choosing one (the preferred) FX relative to the other, which can be interpreted as firms placing higher value on diversification than on fully exploiting perceived arbitrage (carry trade) opportunities, if any, between foreign currencies. Since the outbreak of the crisis, the relative attractiveness of the two main foreign currencies has severely deteriorated. The CHF has lost its relative attractiveness vis-à-vis not only the local currency, but also vis-à-vis the EUR. As a result, a number of firms have switched from the CHF partly to the HUF, but also partly to the EUR.

After briefly introducing the evolution of corporate FX denominated loans in CEE countries and in particular in Hungary (Section 1.2), we discuss the theoretical framework for understanding the trade-offs faced by firms in choosing the currency composition of their debt (Section 1.3). The estimation method is presented in Section 1.4. The dataset used for the estimations is described

in Section 1.5, then the econometric analysis of firms' currency choice is presented in Section 1.6. Finally, Section 1.7 concludes.

## **1.2. Foreign currency debt in Hungary and other CEE countries**

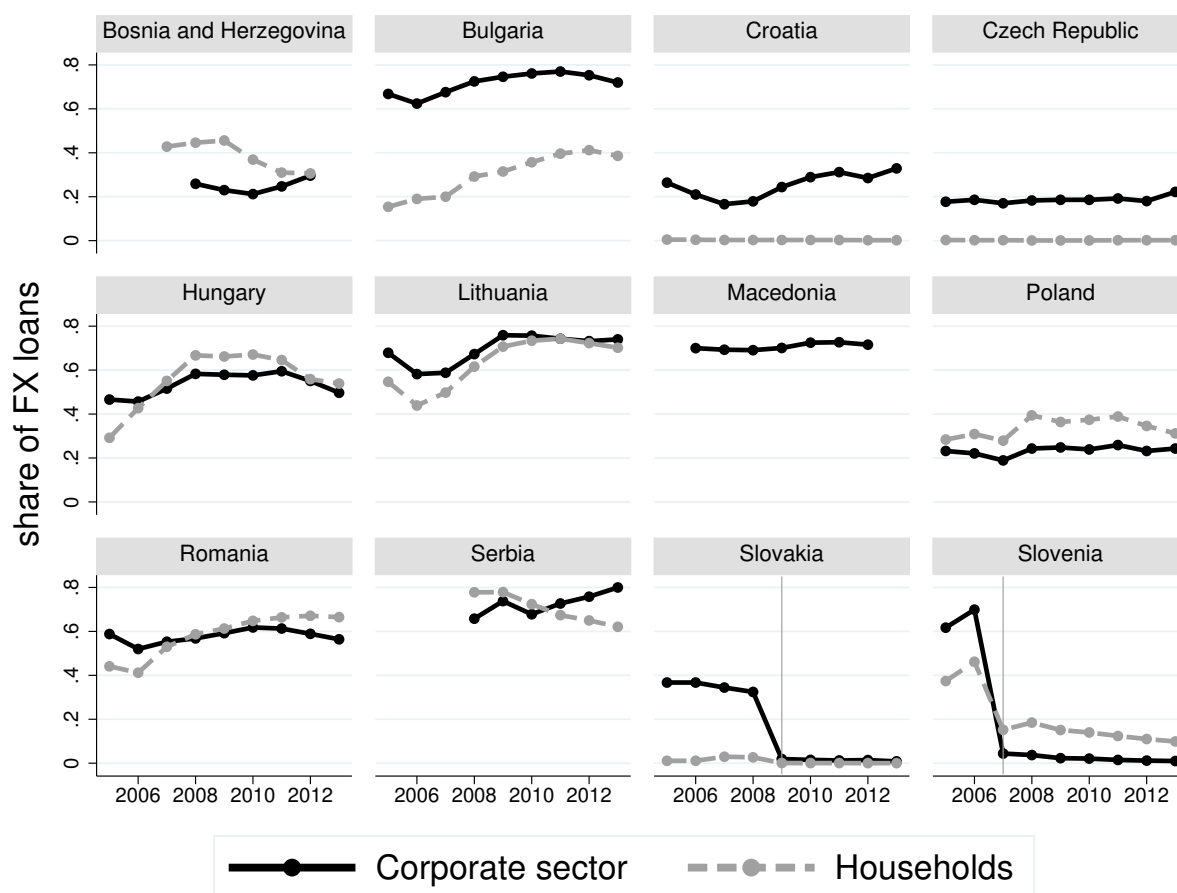
The post-socialist economic transition in emerging European countries was fuelled in part by heavy borrowing from Western banks and easy access to FX denominated loans. For about a decade before the crisis, FX denominated loans became widespread among the majority of Central and Eastern European (CEE) countries. In some of these countries, lending in foreign currencies to both households and firms has been the norm rather than the exception. As shown in Figure 1.1, FX-denominated bank loans accounted for at least 20 percent of total corporate debt in the twelve countries considered. Bulgaria, Macedonia and Slovenia recorded the highest pre-crisis shares, but FX was also dominant in Hungary, Lithuania and Romania. FX lending to the household sector displays a large heterogeneity across countries. It has been mostly prevalent in Hungary, Lithuania, Romania and Serbia, while the share of FX borrowing by households has been virtually zero in Croatia, in the Czech Republic and in Slovakia.

The financial crisis originated from the US subprime mortgage meltdown rapidly escalated to a global scale and brought to the fore the vulnerability of several CEE countries heavily indebted in FX. The high level of bank lending denominated in FX was less of a concern in Slovenia and Slovakia: by the time of the crisis broke out, Slovenia had already adopted the euro, while Slovakia had joined the eurozone at the beginning of 2009. Similarly, the currency board in Bulgaria and the fixed exchange rate regime in Lithuania (and, later, the introduction of the euro) insulated firms and households from adverse exchange rate shocks. In other countries, however, the weakening of the exchange rate first increased firms' debt burden expressed in terms of the local currency then, as a consequence, forced firms to adjust their balance sheets.

Hungary was one of the most affected economies in the region. The country entered the crisis with a combination of a high budget deficit, large current account imbalances and an over-leveraged private sector with a significant exchange rate risk exposure. Already by 2005, more than 45% of the outstanding corporate loans were denominated in FX.



Figure 1.1.: FX shares, 2005-2013

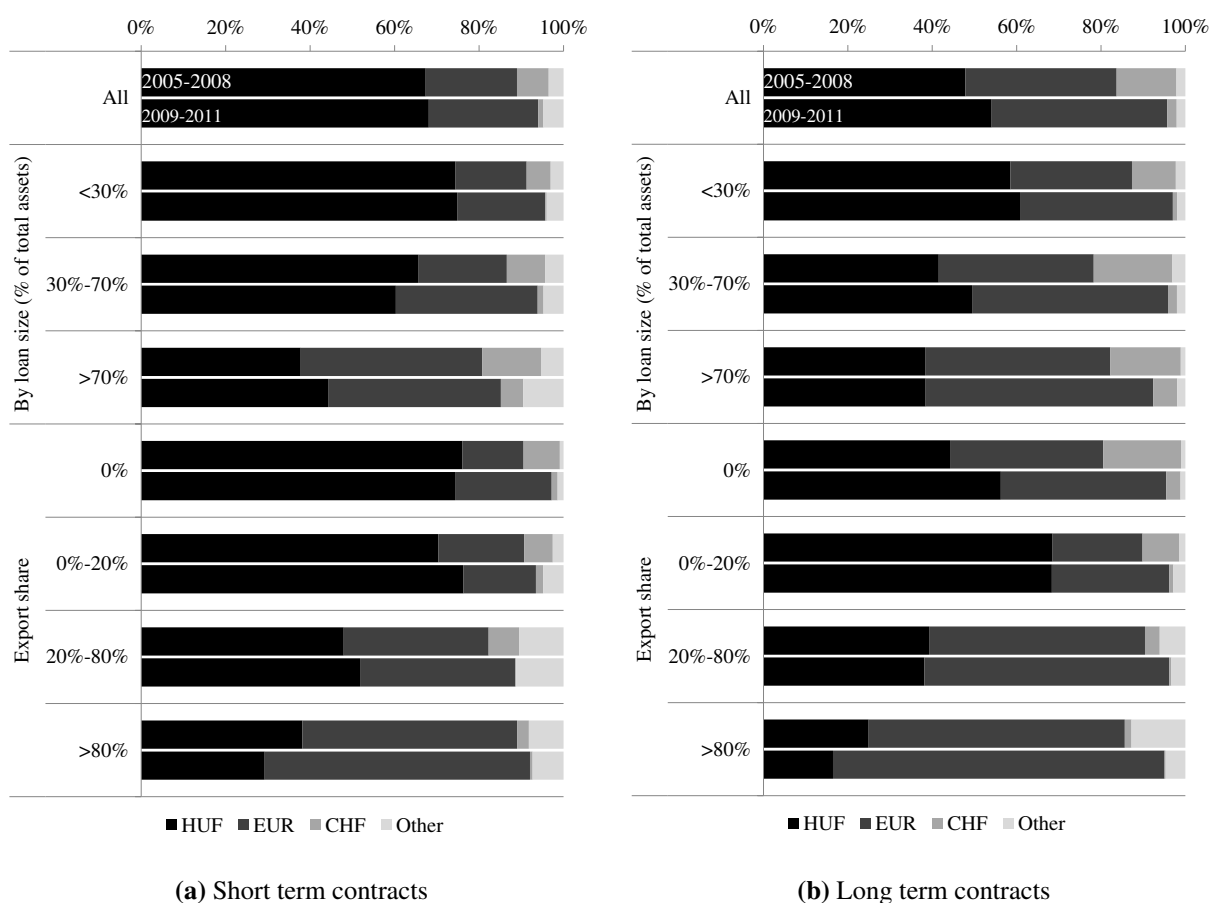


Notes: The figure displays the share of outstanding liabilities in the domestic financial sector (excl. central banks) denominated in foreign currency held by non-financial corporations and households. For Bosnia, Croatia, Macedonia, Slovakia and Slovenia, the data come from the Central Bank's Financial Stability report. For Bulgaria, the Czech Republic, Hungary, Poland and Romania, the data come from the ECB's Balance Sheet Items (BSI) statistics. Vertical lines show the dates of entry into the eurozone: 2007 for Slovenia and 2009 for Slovakia. Lithuania also joined the eurozone in 2015.

Breaking down the composition of debt currencies by maturity and firm characteristics reveal that smaller bank loans – expressed as a percentage of firms' total assets, presumably for continuing operations or for financing replacement investment – were mainly denominated in local currency, while FX loans were primarily used for financing larger projects (Figure 1.2). We arrive at the same conclusion by comparing short-term (less than a year) and long-term contracts separately: comparing subfigures 1.2(a) and 1.2(b) shows that about two-thirds of the overall underwritten sums are in HUF for short-term contracts, but about half for long-term contracts. In line with previous empirical findings (Section 1.1), Figure 1.2 also shows that export-oriented firms are more likely

to take out FX loans than other firms. Moreover, exporters tend to prefer euro-denominated loans. Given that the euro area accounts for an overwhelming share of Hungarian exports, these figures suggest that matching motives are likely to play a role in explaining firms' FX choices. At the same time, the FX debt exposure of non-exporting firms and the relatively large share of CHF loans in firms' debt portfolio also suggest that the share of unhedged loans was (and still is) substantial in Hungary.<sup>2</sup>

**Figure 1.2.:** Characterisation of long and short term contracts



*Notes: The left panel describes short-term loans and a right panel is for long term loans. The Table shows the shares of new loans by denomination. A panel consists of three main blocks. The first describes all loans, with two separate bars for the periods 2005-2011 and 2009-2011, respectively. The second differentiates between loans size relative to the total of the firms securing the loan. We differentiate between three degrees of loan size: below 30%, between 30-70% and above 70% relative to the total assets. The third block differentiates loans by the export share of the firm securing the loan.*

<sup>2</sup>Note that the share of exports to Switzerland in overall exports was only 1.3% in 2008 (source: Eurostat Comext database)

An ongoing research by Vonnak (2014) points to the same conclusion. By comparing CHF and EUR borrowers in the lending boom and during the crisis, the author finds that the latter are more likely to be bigger, foreign-owned and export-oriented firms, while the former are more probably non-exporting firms and firms with weaker balance sheets and a higher default probability already during the pre-crisis period. The descriptive study in Endresz et al. (2012) also reports that FX debt is mostly concentrated among larger and more productive and most likely multinational firms, but a significant share of domestic non-trading firms also took out FX loans. Moreover, a survey conducted by Bodnar (2009) on FX indebtedness of Hungarian companies suggests that financial hedging is practically non-existent. A significant share of the firms simply ignore exchange rate risks as either they are unaware of risk management techniques or such techniques are perceived as expensive, complicated and ineffective. Driven by the attractive foreign interest rates, these firms expose themselves, unwittingly or not, to risks associated with exchange rate depreciation.

Once the crisis broke out, the depreciation of the HUF quickly turned the FX debt previously considered as advantageous into a serious trouble for numerous firms and households heavily indebted in FX. The relative worse post crisis performance of FX borrowers is also confirmed by Endresz and Harasztosi (2014) who demonstrate that FX lending increased investment rates prior to the crisis, while balance sheet effects triggered by the depreciation decreased the investment of firms with FX loans. The authors also show that both effects are likely to be heterogeneous, more pronounced for firms with liquidity constraints. Likewise, Ranciere et al. (2010) show that before the crisis FX borrowing and lower interest rates benefited small domestic firms via a relaxing of credit constraints, while no effect was found for larger firms. The paper does not address firms' post-crisis performance.

After 2008, CHF lending became considerably less attractive. The share of newly contracted bank loans denominated in CHF dropped to one percent by 2010 as compared to ten percent before the outbreak of the crisis. Given the large depreciation of the HUF vis-à-vis the EUR and other hard currencies, it may seem surprising that the share of euro lending did not decrease or even increased among large exporters. There are at least two possible explanations for this observation. First, the composition of firms taking out new loans during the crisis may have changed and the best and most resilient, most probably highly export-oriented firms were more likely to choose euro-denominated loans. Second, firms that previously preferred CHF to EUR may have switched at least partly to euro-denominated loans. These hypotheses will be investigated in detail later in the paper.

Tables 1.1 and 1.2 depict whether the simultaneous presence of several currency denominations at the aggregate level results from the aggregation of distinct individual currency choices or from firms holding multiple currencies in their debt portfolio. The first row of Table 1.1 shows that about one fifth of firms took out both HUF and FX loans during the pre-crisis period. When only firms with more than one loan contract are taken into consideration, this share climbs to 38%. That is, a

## 1.2 Foreign currency debt in Hungary and other CEE countries

large share of firms do not stick to one single currency but choose – presumably strategically – the currency denomination of their loan at all choice occasion. These firms provide useful within-firm variation for our econometric analysis.

**Table 1.1.:** Number of firms by the currency denomination of their contracts (2005-2008)

|   | one contract      |                   | more than one contract |                  |                   | Total   |
|---|-------------------|-------------------|------------------------|------------------|-------------------|---------|
|   | HUF               | FX                | only HUF               | only FX          | both              |         |
| Total number of firms                   | 38 971<br>(36.0%) | 11 032<br>(10.2%) | 33 406<br>(30.8%)      | 2 481<br>(2.3%)  | 22 434<br>(20.7%) | 108 324 |
| only contracts after the first HUF loan | 15 064<br>(28.5%) | 3 339<br>(6.3%)   | 20 480<br>(38.7%)      | 948<br>(1.8%)    | 13 100<br>(24.7%) | 52 931  |
| only contracts after the first FX loan  | 1 035<br>(4.7%)   | 3 320<br>(15.1%)  | 1 104<br>(5.0%)        | 3 436<br>(15.6%) | 13 099<br>(59.6%) | 21 994  |

*Notes: The table shows the number of firms with at least one loan contract underwritten between 2005 and 2008, distributed according to the currency denomination of their loans. The table differentiates between single contract and multi-contract firms. The first row shows all contracts. The second and third rows show only contracts underwritten after the first HUF loan and first FX loan only, respectively. The last two rows thus show data for multi-contract firms only.*

**Table 1.2.:** Number of firms with more than one FX loans (2005-2008)

|                       | only EUR         | only CHF         | only<br>other FX | EUR &<br>CHF     | EUR &<br>other | CHF &<br>other | EUR, CHF<br>& other |
|-----------------------|------------------|------------------|------------------|------------------|----------------|----------------|---------------------|
| two FX loans          | 1 107<br>(18.6%) | 3 978<br>(66.9%) | 42<br>(0.7%)     | 771<br>(13.0%)   | 28<br>(0.5%)   | 18<br>(0.3%)   |                     |
| three FX loans        | 500<br>(21.6%)   | 1 241<br>(53.6%) | 18<br>(0.8%)     | 500<br>(21.6%)   | 23<br>(1.0%)   | 29<br>(1.3%)   | 5<br>(0.2%)         |
| four or more FX loans | 730<br>(24.0%)   | 1 091<br>(35.9%) | 18<br>(0.6%)     | 1 009<br>(33.2%) | 89<br>(2.9%)   | 34<br>(1.1%)   | 67<br>(2.2%)        |

*Notes: The table shows the number of firms with two (first row), three (second row) and four or more FX loans (third row) contracted between 2005 and 2008, distributed according to the currency denomination of their contracts.*

To see whether there is a clear pattern in the order of firms' currency choice, the second and third rows of Table 1.1 display the distribution of firms by their choice of currency denomination of their new loans subscribed after their first domestic or FX denominated loan contract. The third row is particularly interesting, as it disproves the idea that limited access to FX credit is the major source

of within-firm variation. In this latter case, (risk neutral) firms previously indebted solely in HUF which get access to FX loans for the first time during our sample period would never switch back to the local currency. However, more than 69% of firms with an existing FX debt in their balance sheet will later sign at least one HUF-denominated loan contract. In the empirical part of this paper, we will examine whether firms' currency choice and, in particular, the timing of a specific choice, is purely arbitrary or whether it is governed by explicable rationality.

Table 1.2 focuses on firms with more than one FX denominated loan contract. When only firms with two FX denominated contracts are taken into consideration, about 14% of these firms contracted loans in two currencies, mainly EUR and CHF. The share of firms with several foreign currencies in their debt portfolio increases with the number of FX loans contracted: 24% of firms with three FX loans and almost 40% of firms with four or more FX loans prefer to diversify their FX liabilities rather than always choose the same FX.

### 1.3. The optimal debt currency portfolio

To illustrate the effect of exchange rate fluctuations on firms' borrowing decisions, firms' optimal debt currency composition is derived from a simple two-period model. We abstract from some realistic aspects of firms' investment strategies and concentrate only on the main features of their financing decisions that are relevant for our purposes.

The basic structure of our framework is similar to that of Bleakley and Cowan (2008) with two distinctions. First, firms may already hold an initial debt in period 0 when the decision takes place, in which case the amount of remaining debt at the end of period 0 and its interest are reimbursed at the end of period 1. Second, we allow firms to choose – in addition to the local currency – among several foreign currencies in which to borrow.<sup>3</sup> We start with the general case of any arbitrary number of funding currencies, then we look at the simple case with only two possible foreign currencies in more detail.<sup>4</sup>

In period 0, the potential borrowing firm has an initial wealth  $w_0 > 0$  that can be invested and an initial debt  $B_0 \geq 0$  expressed in local currency. A fraction  $\gamma_c$  of the initial debt is denominated in

<sup>3</sup>The model is general regarding the form of debt financing, it applies equally to corporate bonds, commercial papers and bank loans. However, in the empirical part, we will only concentrate on bank loans. The markets for corporate bonds and commercial papers do not exist in Hungary.

<sup>4</sup>As the number of relevant debt issuing currencies is generally limited, firms debt currency allocation problem usually does not require high-dimensional portfolio optimisation. In practice, at most only a few currencies are taken into consideration. One or two currencies generally dominate the FX loan market in all developing and transition countries. In non-European countries, the USD has been the predominant debt issuance currency. Liabilities in emerging Europe have mainly been denominated in EUR and in CHF, the share of other currencies (mainly USD) has been limited.

foreign currency  $c = \{1, \dots, C\}$ . Management decides on the investment strategy, contracts new loans ( $B_1 > 0$ ) to finance the part of the investment that exceeds  $w_0$  and chooses the currency composition of the new loans. The share of the newly contracted debt denominated in foreign currency  $c$  is denoted by  $\beta_c$ . In period 1, the project's cash outflow from the initial investment of amount  $K_1 = B_1 + w_0$  is given by  $F(K_1)$ , where the function  $F$  is assumed strictly increasing and concave in  $K$ .<sup>5</sup> A fraction  $\alpha_c$  of the total output from the project is exported and invoiced in currency  $c$ . Both the previous debt  $B_0$  and the new debt  $B_1$  are reimbursed with the accrued interest at the end of period 1. The firm's terminal net wealth  $w_1$  is given by:

$$w_1 = \left( \sum_{c=1}^C \alpha_c e_{1c} + \left(1 - \sum_{c=1}^C \alpha_c\right) \right) F(K_1) - \left( \sum_{c=1}^C \gamma_c e_{1c} (1 + i_0^c) + \left(1 - \sum_{c=1}^C \gamma_c\right) (1 + i_0^d) \right) B_0 \quad (1.1)$$

$$- \left( \sum_{c=1}^C \beta_c e_{1c} (1 + i_1^c) + \left(1 - \sum_{c=1}^C \beta_c\right) (1 + i_1^d) \right) B_1$$

Today's exchange rates  $e_{0c}$  are normalised to one and we assume that  $E[e_{1c}] = 1$  for all  $c$ . The domestic interest rates of the previous contract and the new contract in local currency are given by  $i_0^d$  and  $i_1^d$ , respectively. The total costs of borrowing in FX ( $i_0^c$  and  $i_1^c$ ) equal the foreign interest rates plus the expected rate of depreciation of the home currency. We assume that the interest rates  $i_0^d$  and  $i_0^c$  are fixed by an already existing contract. For some reason, the UIP between the local currency and any of the foreign currencies may not hold, i.e. the market interest rate in a particular FX may be lower than that of the local currency ( $i^d \geq i^c$ ).

Obviously, any risk-neutral firm would maximise its expected terminal wealth by choosing to incur the full amount of debt  $B_1$  in the "cheapest" currency with the minimum interest rate. In reality, however, the optimal debt portfolio takes into account how investors are averse to risk. The firm's decision is modelled as a mean-variance optimisation problem of modern portfolio theory in which risk averse investors seek to maximise the expected utility of terminal wealth for a given level of risk captured by the variance of the expected cash flows.<sup>6</sup>

For simplicity, we assume that the future exchange rate is the only source of uncertainty, which influences both firms' export revenues and their FX debt reimbursement expressed in local currency.

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<sup>5</sup>Function  $F$  also takes into account capital depreciation.

<sup>6</sup>The mean-variance utility was introduced in the seminal paper by Markowitz (1952), which is considered as the foundation of modern portfolio theory. The concept was later rationalised by Levy and Markowitz (1979), who showed that any twice differentiable von Neumann - Morgenstern utility function can be approximated by a mean-variance utility function. Even though the framework has been relentlessly criticized, the mean-variance technique has a strong intuitive appeal and still constitutes the cornerstone of portfolio theory (Dybvig and Ross (2003)).

From eq. (1.1), the expected variance of the portfolio is given by:

$$\begin{aligned} \mathbf{V}[w_1] = & \sum_{c=1}^C \sum_{c'=1}^C [\alpha_c F(K_1) - \gamma_c(1 + i_0^c)B_0 - \beta_c(1 + i_1^c)B_1] \times \\ & [\alpha_{c'} F(K_1) - \gamma_{c'}(1 + i_0^{c'})B_0 - \beta_{c'}(1 + i_1^{c'})B_1] \rho_{e_c, e_{c'}} \sigma_{e_c} \sigma_{e_{c'}} \end{aligned} \quad (1.2)$$

where  $\sigma_{e_c} > 0$  is the standard deviation of the exchange rate of currency  $c$  vis-à-vis the local currency and  $\rho_{e_c, e_{c'}}$  is the correlation coefficient between the exchange rates of  $c$  and  $c'$ . It is reasonable to assume that  $0 \leq \rho_{e_c, e_{c'}} < 1$  for all  $c' \neq c$ .

According to the mean-variance approach, the firm maximises its utility characterised by the weighted combination of expected terminal wealth and its expected variance:

$$\max_{K_1, \beta} U \left\{ \mathbf{E}[w_1] - \frac{\theta}{2} \mathbf{V}[w_1] \right\} \quad (1.3)$$

where  $\theta > 0$  is the Arrow-Pratt measure of constant absolute risk aversion. The expected terminal wealth is given by eq. (1.1) by setting  $e_{1c} = 1$  for all  $c$  and the expected variance is presented in eq. (1.2).<sup>7</sup>

It is immediately apparent from eq. (1.2) that the minimal variance portfolio is achieved with  $\beta_c = \beta_c^M = (\alpha_c F(K_1) - \gamma_c(1 + i_0^c)B_0) / ((1 + i_1^c)B_1)$  for all  $c$ , i.e. when the currency matching is perfect. In this case, the expected variance of the portfolio is zero. The expected utility with the perfect matching portfolio is a straightforward benchmark that any firm with any degree of risk aversion can achieve. If the firm is not fully risk-averse ( $\theta < \infty$ ) and at least one foreign interest rate  $i_1^c$  is lower than  $i_1^d$ , it is possible to achieve higher expected utility by moving along the efficiency frontier, i.e. choosing a risky (volatile) debt portfolio with higher expected terminal wealth that provides greater utility.

To see this, let us solve the maximisation problem (1.3) w.r.t.  $\beta_c$  for a given (optimal)  $K_1$ . The optimal currency shares are given by the following system of first order equations:

$$\frac{\partial}{\partial \beta_c} = B_1 G_c - \theta B_1^2 \left( \sum_{c'=1}^C (\beta_{c'} - \beta_{c'}^M) \rho_{e_c, e_{c'}} \sigma_c \sigma_{c'} \right) = 0, \forall c \quad (1.4)$$

with  $G_c = \left( (1 + i_1^d) - (1 + i_1^c) \right) \geq 0$  being the expected financial gain from contracting one unit of debt in currency  $c$  instead of the local currency,  $\sigma_c = \sigma_{e_c} (1 + i_1^c)$  is the corresponding standard deviation of the reimbursement obligation.

<sup>7</sup>For presentation convenience, positivity constraints  $-\beta_c \geq 0$  for all  $c$  and  $(1 - \sum_{c=1}^C \beta_c) \geq 0$  are ignored. If any of the constraints are binding, a complementary slackness condition should be applied, which implies a corner solution for one or several currency shares.

Equation (1.4) can be solved using standard linear algebra techniques. Using matrix notations, the solution is given by:

$$\hat{\beta} = (1/(\theta B_1))V^{-1}G + \beta^M \quad (1.5)$$

where  $\hat{\beta}$  and  $\beta^M$  are  $C \times 1$  vectors with elements  $\hat{\beta}_c$  and  $\beta_c^M$ , respectively,  $G$  is an  $C \times 1$  vector of  $G_c$ 's and  $V^{-1}$  is the  $C \times C$  inverse of the variance-covariance matrix (a.k.a. concentration matrix or precision matrix) with elements  $\rho_{e_c, e_{c'}}\sigma_c\sigma_{c'}$ . The optimal portfolio is thus the sum of a standard Markowitz portfolio (the speculative component) and a hedge term represented by the perfect matching portfolio.

It is easy to see that – for a given  $K_1$  – the expected excess financial gain over the perfect matching debt portfolio is  $E[\hat{w}_1] - E[w_1 | \beta = \beta^M] = (1/\theta) G^T V^{-1} G \geq 0$  and the variance of the portfolio is  $V[\hat{w}_1] = (1/\theta^2) G^T V^{-1} G \geq 0$ . That is, if the firm is not fully risk-averse ( $\theta < \infty$ ) and if there is at least one “cheaper” FX in the set of possibilities ( $i_1^c < i_1^d$  for at least one  $c$ ), the firm is willing to take some risk in exchange for higher expected future profits. Indeed, the firm will be taking on a carry trade position by exploiting perceived arbitrage opportunities between funding currencies.

Although the mathematical expressions for  $\hat{\beta}_c$  shares become quite complex as the number of potential currencies increases, the properties of the Markowitz portfolio selection model are well-known from the financial literature. In a world with only one FX, the optimal allocation between the risky FX loan and the risk-free domestic loan is a simple trade-off between the additional gain ( $G$ ) that the firm generates by increasing the FX share above the perfect matching level and the utility lost generated by the higher variance ( $\theta\sigma^2$ ).<sup>8</sup>

The possibility of contracting debt in more than one FX brings in two additional considerations: diversification and possible arbitrage opportunities between foreign currencies. According to equation (1.5), the relative “mismatch shares”  $\tilde{\beta}_c/\tilde{\beta}_{c'} = (\hat{\beta}_c - \beta_c^M)/(\hat{\beta}_{c'} - \beta_{c'}^M)$  for all  $c$  and  $c'$  are independent of the degree of risk aversion and the amount of borrowed funds. That is, firms set their diversification strategies and the relative allocation between foreign currencies according to their beliefs about relative gains and volatilities associated with the various alternatives and the correlation between the exchange rates, independently of their risk preferences.

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<sup>8</sup>According to eq. (1.5), the optimal  $\hat{\beta}$  is also negatively correlated with the amount of borrowed fund  $B_1$ . In fact, in this framework, Markowitz's equations determine the optimal *level* of FX borrowing. It follows that a firm with constant absolute risk aversion ( $\theta$ ) contracts a fixed amount of risky FX debt, independently of the total amount of borrowing and the firm's initial wealth. This unrealistic implication of utility functions with constant absolute risk aversion is largely criticised in the financial literature. Nevertheless, this unpleasant property of the basic mean-variance framework does not alter the main messages of the paper. See e.g. Dybvig and Ross (2003) for alternative utility functions used in the literature.



To probe deeper into how firms choose the relative FX shares of their debt, let us consider the simple case with two available foreign currencies denoted, for example, by *eur* and *chf*. The optimal weights are given by:

$$\begin{cases} \hat{\beta}_{\text{eur}} - \beta_{\text{eur}}^M = \tilde{\beta}_{\text{eur}} = \frac{G_{\text{eur}}}{\theta\sigma_{\text{eur}}^2 B_1} \frac{1 - \rho_{\text{eur}, \text{chf}}(1/\psi)}{1 - \rho_{\text{eur}, \text{chf}}^2} \\ \hat{\beta}_{\text{chf}} - \beta_{\text{chf}}^M = \tilde{\beta}_{\text{chf}} = \frac{G_{\text{chf}}}{\theta\sigma_{\text{chf}}^2 B_1} \frac{1 - \rho_{\text{eur}, \text{chf}}\psi}{1 - \rho_{\text{eur}, \text{chf}}^2} \end{cases} \quad (1.6)$$

The firm's perception of arbitrage opportunities is captured by  $\psi = (G_{\text{eur}}/\sigma_{\text{eur}})/(G_{\text{chf}}/\sigma_{\text{chf}})$ , which is equal to one if the certainty equivalent foreign interest rates (or gains) are equal, and consequently, the firm has no arbitrage incentive between the two foreign currencies.<sup>9</sup> In the absence of arbitrage opportunities, the optimal currency shares simplify to  $\tilde{\beta}_c = (G_c/(\theta\sigma_c^2 B_1))/(1 + \rho_{\text{eur}, \text{chf}})$ ,  $c = \{\text{eur}; \text{chf}\}$ . Both FX shares are strictly decreasing in correlation between the two exchange rates. In fact, the additional FX share above  $\beta_c^M$  is twice as large in the case where the correlation is zero compared to that where  $\rho_{\text{eur}, \text{chf}} \rightarrow 1$ . This result emerges from the principle of Markowitz diversification, which states that as the correlation between the returns on assets that are combined in a portfolio decreases, so does the variance of that portfolio. The same logic applies to optimal debt portfolio choices. The optimising firm can thus increase the share of risky portfolio and therefore increase expected wealth while maintaining risks within acceptable limits.

If  $\psi \neq 1$ , an additional trade-off arises between taking advantage of arbitrage opportunities and diversification. Without loss of generality (as the problem is symmetric), let us assume that  $\psi < 1$ , i.e. *chf* is preferred to *eur*. For low values of  $\rho_{\text{eur}, \text{chf}}$ , the benefit from diversification is relatively high, while the carry-trade between the two foreign currencies is risky. As the correlation between the two exchange rates increases, diversification benefits become smaller and smaller and firms take increasingly advantage of the more attractive currency. Overall,  $\hat{\beta}_{\text{eur}}$  is strictly decreasing with  $\rho_{\text{eur}, \text{chf}}$ , while  $\hat{\beta}_{\text{chf}}$  exhibits a U-shaped relationship with  $\rho_{\text{eur}, \text{chf}}$  with a minimum at  $\rho_{\text{eur}, \text{chf}} = \psi^{-1} (1 - \sqrt{1 - \psi^2})$ .

Although the firm's management is supposed to have a clear preference for *chf* – either because the interest rate and/or the volatility of the exchange rate is lower –, the firm still maintains a higher debt share than required by matching even in the less attractive foreign currency (*eur*) as long as  $\rho_{\text{eur}, \text{chf}} < \psi$ . Above this threshold, the relative attractiveness of *chf* outweighs the diversification benefits. If the positive constraint for  $\beta_{\text{eur}}$  is not binding (i.e. if  $\beta_{\text{eur}}^M > 0$ ), the firm is even willing to sacrifice the security provided by perfect matching and to lower  $\hat{\beta}_{\text{eur}}$  below  $\beta_{\text{eur}}^M$ .

<sup>9</sup>If  $\psi = 1$ , restricting  $\beta_{\text{eur}} = \beta_{\text{eur}}^M$  or  $\beta_{\text{chf}} = \beta_{\text{chf}}^M$  and solving the problem for the other FX share generates, in both cases, the same utility for the firm.

Finally, the full solution of the maximisation problem in eq. (1.3) requires solving for the optimal  $K_1$ . The F.O.C. w.r.t.  $K_1$  yields:

$$F'(K_1) = \left( \sum_{c=1}^C \beta_c (1 + i_1^c) + (1 - \sum_{c=1}^C \beta_c) (1 + i_1^d) \right) + \theta B_1 \left( \sum_{c=1}^C \sum_{c'=1}^C \left( \beta_c - \frac{\alpha_c F'(K_1)}{1 + i_1^c} \right) (\beta_{c'} - \beta_{c'}^M) \rho_{e_c, e_{c'}} \sigma_c \sigma_{c'} \right) \quad (1.7)$$

The optimality condition for  $K_1$  equates the expected marginal product of capital to the user cost (represented by the first term of the right hand side of eq. (1.7)) plus a marginal risk premium (second term). In line with the real option investment theory, the marginal product has to be greater than its marginal cost in the presence of uncertainty (Pindyck (1991)). Uncertainty increases the value of waiting (call option) and decreases the propensity to invest now relative to what would be suggested by a simple net present value rule. In this simple framework, combining equations (1.4) and (1.7) gives:  $F'(K_1) = (1 + i_1^d) / (1 + \sum_{c=1}^C \alpha_c G_c / (1 + i_1^c))$ . Accordingly, the propensity to invest is unaffected by the possibility of borrowing in FX for non-exporting risk-averse investors.

## 1.4. Identification strategy

Using monthly panel data at the contract level, firms' currency choice is studied using discrete choice models. If currency matching is a relevant factor that firms consider in their choice of currency denomination of their bank loans, the probability of contracting new debt in a particular FX must decrease as soon as the firm's debt reimbursement obligation in this currency during a given period of time exceeds its expected export revenues invoiced in the same currency during the same period (see Section 1.3). To test natural matching considerations, we rely on a "mismatch indicator" variable that takes a value of 1 if the firm's debt payment obligations denominated in foreign currency  $c$  (without considering the actual loan contract) is higher than its export revenues in the same foreign currency and 0 otherwise. More precisely:

$$\begin{cases} M_{ijct} = 1 & \text{if } \bar{X}_{ict} - \bar{L}_{ij'c,t+1} < 0 \\ M_{ijct} = 0 & \text{otherwise} \end{cases} \quad (1.8)$$

where  $M_{ijct}$  is the mismatch indicator for firm  $i$  and contract  $j$  in currency  $c$  subscribed in time  $t$ ,  $\bar{X}_{ict}$  denote the firm's past 12-months average export revenues invoiced in currency  $c$  and  $\bar{L}_{ij'c,t+1}$  is the firm's monthly average debt payment obligation over the next 12 months in the same currency  $c$  stemming from all existing contracts  $j'$  other than the actual loan contract  $j$ . We only use past

values for  $\bar{X}_{ict}$  – just as for all other explanatory variables explained later – to avoid simultaneity. The mismatch indicator is thus equal to 1 if the firm is already in mismatch without considering the actual loan contract.

In our empirical specifications, firms' decisions are modelled as a probabilistic choice problem. In line with the theoretical predictions of optimal currency shares presented in Section 1.3, the extent of deviation from the perfect matching portfolio is influenced by firms' risk perceptions, risk attitudes and expectations about the future paths of interest rates and exchange rates, all of which are subjective assessments that investors believe in and that differ from firm to firm. These measures are captured by a set of observed time varying firm-level and bank-level characteristics ( $Z_{it-1}^T$ ), a currency-specific firm-level unobserved parameter ( $a_{ic}$ ) that represents the effects of firms' unobserved attributes and a random component  $\varepsilon_{ijct}$  assumed to follow an i.i.d. logistic distribution. Under these assumptions, the conditional probability that currency  $c$  is chosen is given by (see McFadden (1974)):

$$P(y_{ijt} = c \mid Z_{it-1}^T, M_{ijct}, a_{ic}, \forall c) = \frac{\exp(a_{ic} + Z_{it-1}^T \Omega_c + \phi_c M_{ijct})}{1 + \sum_{c'=1}^C \exp(a_{ic'} + Z_{it-1}^T \Omega_{c'} + \phi_{c'} M_{ijc't})} \quad (1.9)$$

where  $y_{ijt}$  is the observed outcome.<sup>10</sup> Parameters  $\phi_c$  capture firms' willingness to match the currency composition of their incomes and liabilities to avoid exposure to exchange rate risk. If matching incentives matter, we expect that a firm is less likely to take out a FX loan in a situation of currency mismatch, so  $\phi_c$  is expected to be negative. The baseline category is the local currency (HUF) with the probability of being chosen given by  $P(y_{ijt} = \text{HUF} \mid Z_{it-1}^T, M_{ijct}, a_{ic}, \forall c) = 1/(1 + \sum_{c'=1}^C \exp(a_{ic'} + Z_{it-1}^T \Omega_{c'} + \phi_{c'} M_{ijc't}))$ .

### 1.4.1. The binomial case

In our first specification, we estimate the impact of the mismatch indicator on the probability of signing a FX loan contract rather than one in local currency. For each month, all foreign currencies are collapsed together. The set of possible alternative choices is thus limited to  $c = \{\text{fx}\}$  and HUF

<sup>10</sup>McFadden derives the analytical expression for the selection probabilities in eq. (1.9) using the axiom of independence of irrelevant alternatives (IIA) introduced by Luce (1959), which states that the relative odds of one alternative being chosen over a second one is independent of the presence or absence of any other alternatives. Under this assumption, the relative odds of choosing a specific foreign currency rather than the local currency can be determined as if no other foreign currency alternative were available. Accordingly, the probability of choosing foreign currency  $c$  is given by  $P[(U \mid \beta_c = 1) > (U \mid \beta_c = 0)] = P[G_c/(\theta \sigma_c^2 B_1) + \beta_c^M > 1/2]$  or, equivalently:  $P[a_{ic} + Z_{it}^T \Omega_c + \phi_c M_{ijct} > \varepsilon_{ijct}]$ . With multiple foreign currencies, the system of independent logit equations leads to the expression for the probability that firm  $i$  chooses currency  $c$  given by eq. (1.9). As explained later, the strong assumption of IIA can be relaxed by specifying, for example, a mixed logit model.

as the baseline category. The matching indicator is constructed using total export revenues and total debt payment obligations all foreign currencies combined. In this case, equation (1.9) reduces to the binomial logistic function (and subscripts  $c$  can be dropped).

The main econometric difficulty is dealing with unobserved heterogeneity ( $a_i$ ), in particular its relationship with the covariates. Explicitly including dummies for the fixed effects and estimating the equation using standard logit yields consistent estimates only if the time dimension tends to infinity. For fixed time dimension ( $\Gamma$ ), the unconditional maximum-likelihood estimator of the incidental parameters is inconsistent, which in turn contaminates the rest of the coefficients. The inconsistency arises because the number of incidental parameters increases without bound, while the amount of information about each incidental parameter remains fixed (Neyman and Scott (1948)). To resolve the endogeneity issue due to the presence of incidental parameters, Andersen (1970) and Chamberlain (1980) propose an estimator of the structural parameters by conditioning the likelihood function on minimal sufficient statistics for the incidental parameters and then maximizing the conditional likelihood function. In the logit case, such statistics can be  $\sum_{t=1}^{\Gamma} y_{ijt}$ . Intuitively, the minimal sufficient statistics capture all possible information about time-invariant firm-level parameters which influence how many times an alternative has been chosen by the firm. Conditional on this, the parameters of interest are identified by using information on when a specific alternative is chosen.

The principal advantage of Chamberlain's conditional (fixed effects) logit is that it requires no assumption on  $a_i$ , hence it allows for any form of correlation between the fixed effects and the regressors. However, the estimation method has several drawbacks. First, since the parameters are identified using within-firm variation, only firms which change state (i.e. those indebted in more than one currency) are considered. Although the sample of firms with bank loans denominated in more than one currency is large enough for asymptotic results to be valid (see Table 1.1), the incomplete coverage of firms might be a problem if one wants to draw inferences for the whole population or the excluded sub-population. Second, the incidental parameters are not identified and their distributions are unrestricted, which are necessary to calculate quantities of interest such as marginal effects and probability projections. Finally, the conditional logit exhibits the unpleasant property of independence of irrelevant alternatives (IIA): adding another alternative (another foreign currency, in our case) does not affect the relative odds between the two alternatives previously considered.<sup>11</sup>

An alternative approach is to treat the unobserved heterogeneity as random effects. Obviously, the extreme assumption of no-correlation between  $a_i$  and the covariates is necessarily violated.

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<sup>11</sup>See e.g. Wooldridge (2010). This property is irrelevant if only two options are taken into consideration: foreign currency and local currency. However, in the multinomial case, the model generates implausible substitution patterns.

Indeed, a more risk-averse firm is, *ceteris paribus*, less likely to carry FX debt and its export revenues are thus more likely to exceed its FX debt payment obligations. Mundlak (1978) and Chamberlain (1982) relax this crucial random effects assumption by allowing the unobserved effects to be correlated with the covariates following a linear specification. In Mundlak's specification,  $a_i = \bar{X}_i^T \xi + \omega_i$ , with  $\bar{X}_i^T$  being a row vector of the time-average of all exogenous covariates ( $Z_{it-1}^T$  and  $M_{ijt}$ ) and  $\omega_i$  being a normally distributed error term. Chamberlain proposes a more general form by including the vector of all explanatory variables across all time periods:  $a_i = \sum_{t=1}^T X_{it}^T \xi_t + \omega_i$ . In both cases, the additional explanatory variables included in the model allow us to control for the correlation between  $a_i$  and the covariates while using a standard random effects estimator. The intuition behind the identification is in fact similar to that of the conditional logit estimator. In this paper, we employ both the conditional logit model and the correlated random effects logit model with Mundlak's correction to estimate the determinants of foreign currency choice.

### 1.4.2. The multinomial case

To estimate the discrete choice model with all available alternatives, a separate mismatch indicator has to be constructed for all foreign currencies. Unfortunately, export destinations (or, which would be even better, export invoicing currencies) are not specified in the database that we use. Based on the relative importance of the different currencies in the Hungarian external trade and the aggregate currency composition of the loans, it is reasonable to assume that the main "matching currency" that firms may consider to hedge exchange rate risks on exports is EUR, while CHF is the principal "speculative currency" irrelevant for hedging purposes. Indeed, the euro area is Hungary's major trading partner: it accounts for more than 57% (in 2008) of the country's total exports, compared to less than 2% for Switzerland. We therefore rely on the simplifying assumption that all export revenues are invoiced in EUR for all firms. The other foreign currencies are collapsed together and we estimate a three-alternative choice model with  $c = \{\text{eur ; other foreign currencies}\}$  and HUF as the baseline category. The mismatch indicator is constructed for the EUR only. As an alternative, we also test the model which places emphasis on the CHF as a speculative currency. The choice set becomes:  $c = \{\text{chf ; other foreign currencies}\}$  and HUF. Since the share of foreign currencies other than the EUR and the CHF has been limited in Hungary, we expect that the two specifications yield similar results. For presentation convenience, in what follows, we simply denote the set of choices by  $c = \{\text{eur ; chf}\}$ .

The multinomial discrete choice models are estimated using the mixed logit procedure described in detail in Train (2003). The approach allows very flexible substitution patterns through the estimation of random rather than fixed parameters. Within each firm, the random terms are allowed to be

correlated across alternatives, however, they are uncorrelated across firms:  $\text{corr}(\omega_{ic}, \omega_{i'c'}) = \rho_{\omega_c, \omega_{c'}}$  for all  $c$  and  $c'$  if  $i = i'$  and 0 otherwise.

Conditional on knowing the parameters of the model, the probability that firm  $i$  chooses currency  $c$  on a given choice occasion is given by McFadden's logit formula (eq. 1.9). The unconditional mixed logit probability is the integral of the conditional probability over all the possible parameter values, which depends on the density function of each of the random parameters. The estimation is carried out using maximum simulated likelihood technique (see Train (2003)).

Like in the binomial case, we apply Mundlak's correction to both alternatives by including firm level averages in the equations. Theoretically, it is possible to use a very general random coefficient specification by assuming all coefficients ( $\Omega_c, \phi_c, \xi_c$  and the intercepts for all  $c$ ) to vary randomly. However, depending on the number of coefficients the estimation procedure becomes very complex as multiple integrals have to be solved (Train (2003)). In this paper, we only assume the intercepts to be random.<sup>12</sup>

## 1.5. The dataset

Estimations were carried out using four matched administrative datasets. We principally rely on the credit register (*KHR*) database containing the universe of all new and already existing corporate loan contracts from Hungarian financial institutions between 2005 and 2011. The dataset provides information on contracts starting date, duration, value, denomination, loan types and type of providers. Four types of contracts are distinguished: loans, credit lines, factoring and leasing. Providers are banks, saving banks and other financial companies. Between 2005 to 2011, the average annual number of new contracts stands at 65,000. It rises from about 77,000 to a little over 82,000 until the outbreak of the crisis. After 2008, it drops to below 50,000. Overall, there are 129,066 firms in the dataset and, on average, about 42,700 firms take out loans each year.

For the detailed description of the *KHR* dataset, see Endresz et al. (2012). We deviate from the construction of the dataset described in Endresz et al. (2012) on one important aspect. To focus on currency choices, we collapse the loan contacts denominated in the same currency and signed by the same company in the same month. That is, if a firm takes out two loans in the same currency in the same month, we combine the corresponding contracts to form a single contract with the sum of the two loans and a duration defined as the weighted mean of the duration of the two original loans.

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<sup>12</sup>The mixed logit procedure with an arbitrary combination of random and fixed parameters can also be applied to the binomial case. Assuming only the intercept to be random is equivalent to the random effects logit model described earlier. Although the estimation strategy of mixed logit models differs from that of the standard random effects model, the mixed logit procedure with random intercepts and the random effects logit model yield similar results.

As a result, the average annual number of new contracts falls to about 70% of the original, while the total amount of outstanding debt in each month and the aggregate monthly flow of debt service expenses remain unchanged.

The credit register is then merged with the yearly panel database of corporate tax returns. The database is provided by the Hungarian tax authorities (*NAV*) and contains balance-sheet and income statement information for all double entry book-keeping firms operating in Hungary. We use variables that are likely to affect firms' demand for credit and its choice of currency denomination such as employment, foreign ownership, capital, liquidity, total assets and profitability measures.

Although the *NAV* dataset contains the export share of sales, we collect additional trade information from the Hungarian Central Statistical Office on trade behaviour. We merge the statistics on exports and imports calculated from the monthly reports on commodity trade to Extra- and Intrastat for the universe of direct trading firms in Hungary. The monthly frequencies enable us to calculate, for example, the export revenues during the 12 months preceding the signature of the loan contract.

Finally, we extend the dataset by including information on the credit provider. Following the methodology proposed by Ongena et al. (2014), we match a list of bank characteristics – such as foreign ownership, total assets, capital ratio, liquidity ratio, return on assets and doubtful loan ratio – to the contract-firm level dataset.<sup>13</sup>

## 1.6. Estimation results

### 1.6.1. Pre-crisis period

Table 1.3 summarises the main empirical findings for the pre-crisis period (2005-2008). The first two columns report the results for the binomial specification estimating the probability of choosing a foreign currency over the domestic currency using conditional (fixed effects) logit (column (1)) and Mundlak's correlated random effects logit (column (2)) regression techniques. In both cases, the coefficients of the mismatch indicator are negative and highly significant. The probability of taking out an FX loan is thus higher as long as the firm's expected export revenues fully cover its foreign currency debt service expenses, which provides strong evidence to support the role of natural hedging incentives in firms' debt currency choice.

The mixed logit models with three possible alternatives provides a more accurate, albeit still imperfect, measure of currency matching. As explained in Section 1.4, two specifications are considered.

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<sup>13</sup>We thank Adam Szeidl for giving us access to the structured dataset and Dzsamila Vonnák for the excellent work of matching the databases.

## 1.6 Estimation results

First, the results presented in column (3) of Table 1.3 correspond to the case where exports are all assumed to be invoiced in euro and consequently, only the euro-denominated debt is used for hedging purposes. Debt incurred in other foreign currencies (that are collapsed together) is the result of pure speculation. Second, column (4) reports the results for the case in which CHF is the only purely speculative currency and debt incurred in any other foreign currency may potentially be used for hedging purposes. The mismatch indicators are constructed accordingly.

**Table 1.3.:** Estimation results for the pre-crisis period (2005-2008)

|   | Fixed-effects<br>logit | Random-<br>effects<br>logit | Mixed logit with three alternatives                             |   |
|---|------------------------|-----------------------------|---|---|
|   |                        |                             | EUR, other<br>foreign<br>currencies<br>( $\sim$ CHF) and<br>HUF | CHF, other<br>foreign<br>currencies<br>( $\sim$ EUR) and<br>HUF |
|   | (1)                    | (2)                         | (3)   | (4)   |
| Mismatch indicator                                      | -1.058***<br>[0.025]   | -1.244***<br>[0.027]        | -0.987***<br>[0.055]  | -0.923***<br>[0.054]  |
| $\text{var}(\omega_{\text{eur}})$                       |                        |                             | 3.087***<br>[0.112]   | 3.152***<br>[0.113]   |
| $\text{cov}(\omega_{\text{eur}}, \omega_{\text{chf}})$  |                        |                             | 0.553***<br>[0.090]   | 0.413***<br>[0.0573]  |
| $\text{var}(\omega_{\text{chf}})$                       |                        |                             | 2.443***<br>[0.055]   | 2.465***<br>[0.055]   |
| $\text{corr}(\omega_{\text{eur}}, \omega_{\text{chf}})$ |                        |                             | 0.201***<br>[0.032]   | 0.148***<br>[0.020]   |
| Nb. of obs.   | 83 147                 | 173 673                     | 521 019   | 521 019   |
| Nb. of firms  | 12 869                 | 51 706                      | 51 706  | 51 706  |

*Notes: Each column of Table 1.3 collects results from a separate regression. The first specification is a firm fixed effects logit. It includes all firm- and bank-level controls (both lagged) described in the Appendix, year dummies and a constant. The next column shows the results for the correlated random effects logit regression. In addition to the aforementioned controls, the regression also controls for the time-averages of the variables for each firm to implement Mundlak's correction. The last two columns collect results from mixed logit regressions with three alternatives. Here, we allow the constants in both choice equations to be randomly distributed across firms. The elements of the covariance matrix and the correlation between alternatives (i.e. between the random terms  $\omega_{\text{eur}}$  and  $\omega_{\text{chf}}$ ) are calculated from the estimated lower-triangular matrix L, where the matrix L is the Cholesky factorization of the covariance matrix. The corresponding standard errors are computed using the delta method. \*\*\* significant at 1%, \*\* significant at 5%, \* significant at 10%.*



The two specifications yield similar results. In particular, both estimates reinforce the role of matching incentives in firms' currency choice and suggest that, as expected, results previously presented for the binomial case are mainly driven by matching in euro.

The positive and significant estimated covariance between the two foreign currency alternatives implies that firms that are more likely to choose one of the two foreign currencies are also more likely to choose the other. Theory (Section 1.3) suggests that the substitution pattern between foreign currencies – i.e. the within-firm correlation between the random effects – is affected by two distinct yet interacting factors: arbitrage and diversification. While our model does not allow us to disentangle the two factors, results suggest that the benefits from diversification – i.e. holding both EUR and CHF debt – outweighs the advantage of consistently choosing one (the preferred) FX to the other. In other words, the average (or representative) firm seeks diversification. However, the correlation coefficient of about 0.2 implied by the estimated parameters is rather low. As the average number of contracted FX loans is relatively low in our sample (1.8 among the firms that subscribed at least one FX contract; see also Table 1.2), the model fails to detect a strong interdependence between the alternative choices.

We also check the robustness of our results to various sample selection alternatives. The estimation sample for the conditional fixed effects model is by construction restricted to firms indebted in more than one currency. In principle, however, all observations can be used to estimate the correlated random effects and the mixed effects models, even though the firms that exhibit no variation in the explanatory variables or the dependent variable do not contribute to the identification of the structural parameters. Along with our baseline specification on the whole sample of firms, we re-estimate our models presented in columns (2) to (4) of Table 1.3 for the sample of firms exhibiting a variation in the left-hand side variable (same sample as for the fixed effects logit) and for the sample of firms with variation in the matching indicator. In addition, we test whether using net exports instead of export sales or considering only long-term contracts changes the results of the fixed effects, random effects or mixed logit models. Finally, we re-estimate all our models separately for foreign firms (foreign ownership over 50%) and domestic firms (we have run the regressions both on the sample of firms with foreign ownership less than 50% and on the sample of entirely domestic firms). For all estimation methods, results based on these alternative sample selection choices are very close to our baseline findings.<sup>14</sup>

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<sup>14</sup>These results are not presented in this paper for reasons of brevity, but are available from the authors upon request.

### 1.6.2. The effects of the crisis on firms' currency choice

Our final specification estimates the mixed logit model in column (3) of Table 1.3 for the whole period covered by our database (2005-2011). To explore potential changes in firms' matching behaviour, in particular after the outbreak of the crisis, we let the parameter of the matching indicator take different values for the pre- and post-crisis periods. As explained in Section 1.4, different random parameters ( $\omega_{ic}$ ) are estimated for the pre- and post-crisis periods that can be correlated both in time and between alternatives. The estimation sample covers only firms contracting new debt in both periods. Figure 1.3 plots the evolution of the estimated average alternative-specific effects calculated from the year dummies and the mean estimates of the random parameters.<sup>15</sup>

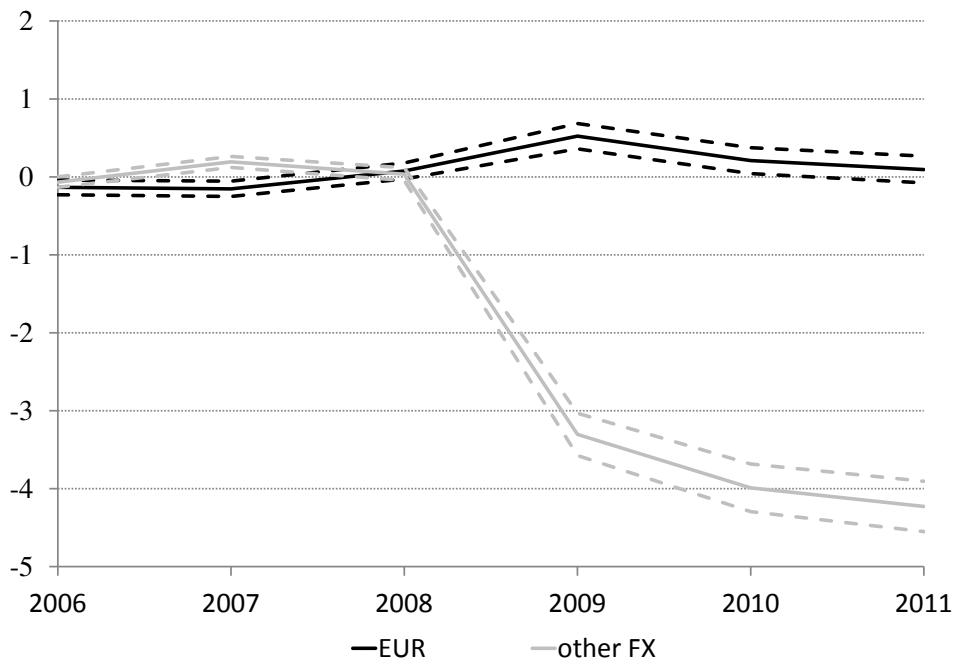
The significantly negative interaction term between the mismatch indicator and the post-crisis dummy suggests that matching incentives played a somewhat more important role in firms' currency choice decisions in the aftermath of the crisis than before 2008.<sup>16</sup> At the same time, the estimated average year effects (Figure 1.3) reveal that the attractiveness of the two foreign currencies remained stable prior to the crisis but has severely deteriorated since 2009. While the relative odds of taking out euro-denominated corporate loans compared to local currency loans did not change significantly in the aftermath of the financial crisis, either the expected financial gain from taking out CHF loans declined or the perceived risks associated with bank loans denominated in CHF increased considerably.<sup>17</sup>

<sup>15</sup>More precisely, the figure plots the year dummy parameters ( $\hat{d}_y$ ) for the years 2006 to 2008. For 2009, it is calculated as the difference between the mean of the random parameters:  $\hat{d}_{2009} = \hat{\omega}_{c,\text{post}} - \hat{\omega}_{c,\text{pre}}$ . For the years 2010 and 2011, we subtract  $\hat{d}_{2009}$  from the estimated year dummies.

<sup>16</sup>The point estimate for the parameter of the mismatch indicator for the pre-crisis period is lower in absolute value than the parameter values presented in Table 1.3. These estimates are, however, not directly comparable. As is the case with any nonlinear probability model, the coefficients are identified only up to scale and therefore estimated parameters across various estimation methods and different sample of firms and periods cannot be naively compared.

<sup>17</sup>We interpret estimates for "other foreign currencies" as results for CHF for two reasons. First, as explained earlier in the paper, the share of foreign currency loans other than EUR and CHF has been limited in Hungary and consequently, contracts denominated in CHF may drive the results. Second, our alternative mixed logit specification presented in column (4) of Table 1.3 re-estimated for the whole period 2005-2011 with time-varying matching indicator parameters yield similar results to the specification discussed in the paper.



**Figure 1.3.:** Average currency-specific effects

*Notes: The Figure plots the evolution of the estimated average alternative-specific effects issued from the mixed logit regression on the whole sample 2005-2011. More precisely, the values plotted on the graph are: year dummy parameters ( $\hat{d}_y$ ) for the years 2006-2008; for 2009, it is calculated as the difference between the mean of the random parameters  $\hat{d}_{2009} = \hat{\omega}_{c,post} - \hat{\omega}_{c,pre}$ ; and from the years 2010 and 2011,  $\hat{d}_{2009}$  is subtracted from the estimated year dummies  $\hat{d}_{2010}$  and  $\hat{d}_{2011}$ , respectively. The standard errors corresponding to the calculated parameters are computed using the delta method.*

The estimated covariance matrix of the random coefficients is also consistent with the deterioration in firms' relative preference for the EUR over the CHF. The relatively high correlations between pre- and post-crisis random terms for both currencies – over 0.7 for both the EUR and the CHF, see Table 1.4 – indicate that firms that placed a higher value on a particular currency before the crisis still have a stronger preference for this currency than the average firm. However,  $\text{cov}(\omega_{\text{chf, pre-crisis}}, \omega_{\text{eur, post-crisis}})$  largely exceeds  $\text{cov}(\omega_{\text{eur, pre-crisis}}, \omega_{\text{chf, post-crisis}})$ , which suggests that the probability of firms with a stronger preference for the CHF before the outbreak of the crisis switching to the EUR is greater than the probability of firms with a stronger preference for the EUR before the crisis of later switching to the CHF. In other words, a number of firms seem to have adjusted their optimal relative shares of the two FX-denominated debts and have switched from the CHF partly to the HUF, but also partly to the EUR.

### 1.6.3. Matching or speculation?

An important advantage of our modelling approach compared to those previously used in the literature is that our model allows us to perform a counterfactual analysis to isolate the effects of currency matching motives on the aggregate corporate FX debt share. To do so, we first predict for all firms and for all choice occasions the probability of choosing the EUR, the CHF or the HUF using the estimated model of Table 1.4. The weighted – by the size of the loan – yearly averages of these probabilities give the estimated aggregate currency shares of newly contracted bank loans. If the fit of our model is good, the estimated shares should be close to the observed shares. We then “switch off” the effects of currency matching on firms’ debt denomination choice by setting the mismatch indicator to 1 for all observations and we predict the counterfactual currency shares of newly contracted corporate loans in the absence of matching motives. The weighted averages of these counterfactual probabilities correspond to the currency shares of new loans that would have resulted from “speculation”, i.e. if all firms were constantly in a mismatch situation and hedging strategy was irrelevant. The difference between the baseline predictions and the counterfactual currency shares represents the effect of matching on the aggregate shares.

**Table 1.5.:** Decomposition of currency choice

|      | Prediction |      |       |          |      |       |                    |      |       |                 |     |       |
|------|------------|------|-------|----------|------|-------|--------------------|------|-------|-----------------|-----|-------|
|      | Data       |      |       | Baseline |      |       | “Speculative” part |      |       | Matching effect |     |       |
|      | (1)        |      |       | (2)      |      |       | (3)                |      |       | (4) = (2) - (3) |     |       |
|      | HUF        | EUR  | other | HUF      | EUR  | other | HUF                | EUR  | other | HUF             | EUR | other |
| 2005 | 55.0       | 31.7 | 13.3  | 55.7     | 33.6 | 10.7  | 60.3               | 28.2 | 11.5  | -4.6            | 5.4 | -0.8  |
| 2006 | 53.7       | 32.0 | 14.2  | 58.5     | 31.0 | 10.5  | 63.6               | 25.1 | 11.4  | -5.1            | 5.9 | -0.9  |
| 2007 | 54.2       | 29.7 | 16.0  | 58.8     | 27.9 | 13.2  | 63.2               | 22.6 | 14.2  | -4.4            | 5.3 | -1.0  |
| 2008 | 60.8       | 26.4 | 12.7  | 61.5     | 27.2 | 11.3  | 65.8               | 22.1 | 12.1  | -4.3            | 5.1 | -0.8  |
| 2009 | 62.8       | 32.1 | 5.1   | 60.1     | 38.2 | 1.8   | 68.4               | 29.6 | 2.0   | -8.3            | 8.6 | -0.2  |
| 2010 | 62.0       | 33.7 | 4.3   | 62.0     | 36.9 | 1.1   | 69.3               | 29.4 | 1.2   | -7.3            | 7.5 | -0.1  |
| 2011 | 62.4       | 30.9 | 6.7   | 63.4     | 35.7 | 0.9   | 71.0               | 28.0 | 1.0   | -7.6            | 7.7 | -0.1  |

*Notes: The first block displays the currency shares of newly contracted loans in the database between 2005 and 2011. The second block shows the weighted averages of the predicted currency shares using the estimated mixed logit model of Table 1.4. The “speculative part” in column (3) is obtained by “switching off” the matching motive in firm’s decision, i.e. by setting the mismatch indicator to one for all firms and all choice occasion and predicting the currency shares the same way as before. The effects of the matching, presented in the last block of the Table, is the difference between the baseline predictions and the predicted speculative part.*

Table 1.5 presents the results. The first block of the table displays the observed currency shares of newly contracted loans in the database between 2005 and 2011, while the second block shows the baseline predicted currency shares using the estimated mixed logit model of Table 1.4. Although the model appears to moderately underestimate the share of “other FX” ( $\sim$ CHF), overall the predicted currency shares are close to the observed ones.

If currency matching had not been part of firms’ strategies, the share of the newly contracted FX debt would have been lower by 4.6 percentage points before the crisis and by about 7.7 percentage points after the outbreak of the current crisis (see blocks (3) and (4) of Table 1.5). Given that the share of the new FX loans is on average approximately 41 percent between 2005 and 2008 and 38 percent between 2009 and 2011, roughly 10 percent of the overall corporate FX debt during the pre-crisis and 20 percent during the post-crisis periods can be explained by matching motivation. The largest share of FX debt is thus not related to natural hedging.

## 1.7. Conclusion

Households indebted in FX naturally expose themselves to exchange rate risks. Conversely, firms with export revenues can also use FX debt to reduce or eliminate their exposure to exchange rate variations. While a large body of empirical research documents a significant positive correlation between the share of FX debt in firms’ balance sheets and a proxy for the sensitivity of firms’ revenues to exchange rate fluctuations, such as export shares or an indicator of tradability, there is little rigorous empirical evidence on the importance of matching motives in firms’ currency-of-denomination decisions. Furthermore, there has been no clear understanding whether natural hedging motivation or other factors such as the interest rate differential is the key driver of firms’ currency choices.

This paper investigates firms’ willingness to match the currency composition of their assets and liabilities and their incentives to deviate from perfect matching in the presence of multiple available foreign currency loans. By adopting a simple mean-variance approach from modern portfolio theory, we first derive a closed form solution for optimal debt portfolio. In line with the proposed theory, we then rely on Hungarian corporate loan data and estimate discrete choice models in which firms choose the currency denomination of their loans.

Results show that the probability of borrowing in FX decreases as soon as the firm’s foreign currency debt reimbursement obligation exceeds its expected export revenues. This finding is robust across various model specifications and sample choices, which provides strong evidence to support the role of currency matching incentives in firms’ currency choice. In addition, our results suggest

that the benefits from diversification outweigh the perceived carry trade opportunities between EUR and CHF, the two major foreign currencies in Hungary.

Matching motivation is even stronger after the outbreak of the current financial crisis than during the pre-crisis period. However, our counterfactual simulations suggest that natural hedging is not the main cause of firms' FX indebtedness: only about 10 percent of overall corporate FX debt is attributable to natural hedging during the pre-crisis and 20 percent during the post-crisis periods.

While our model allows us to isolate the effects of matching motives on firms' currency choice, we do not directly explain the underlying reasons why firms deviate from the pure natural hedging strategy. One obvious candidate is the interest rate differential and thus firms' carry trade strategies, yet other explanations may also exist. The existing literature proposes a few other explanations for firms' FX choice. However, none of these alternative explanations seem relevant for Hungary. For instance, Shapiro (1984) shows that in some countries, such as Sweden, the tax law encourages firms to incur FX debt by making foreign exchange losses on FX debt immediately tax deductible, while taxes on foreign exchange gains are deferred until realised. On the other hand, if the exchange losses (gains) on the principal of the FX debt are tax deductible (taxable) at the same rate as the local corporate income tax, such as is the case in Hungary, the firm is indifferent between borrowing in FX or local currency. Other papers concentrate on the financing decisions of multinational firms. Multinationals may have incentives to locate their debt in the highest-tax country (see e.g. Hodder and Senbet (1990)) or may choose the country and the currency in which the debt is incurred depending on legal barriers (Jorion and Schwartz (1986)) or the costs of gathering information (Hietala (1989)). These theories can hardly explain the FX borrowing of many domestic firms from local commercial banks.

Other explanations may exist, but apart from exploiting interest rate differentials, we are not aware of any convincing evidence or theory that could explain the large share of FX debt in Hungary or in other similar countries. Most likely, the largest share of corporate FX debt, at least in Hungary, corresponds to open carry trade positions held by non-financial corporations.

## 2. Income Taxation, Transfers and Labour Supply at the Extensive Margin

### 2.1. Introduction

This paper proposes the “gains to work” model to estimate the impact of taxes and transfers on the static participation decision, i.e. the labour supply at the extensive margin. The initial setup is similar to the discrete choice approach based on random utility models (van Soest (1995)), in which utility-maximising individuals choose between discrete sets of hours worked based on their consumption-leisure preferences. However, instead of explicitly estimating the structural parameters of the full underlying preference specification, we derive – from a standard and general utility function – an expression for a “reservation gains to work”, i.e. the minimal difference between net wages and the amount of lost transfers required by the individual to accept a job offer. The reservation gains to work then translates into a labour supply function in which the individual’s choice depend on the financial gains to work, the non-labour income and other individual characteristics. The basic underlying concept of the model is similar to the labour supply models based on marginal calculus (Hausman (1981)), but it is adopted to reflect the nonlinearities of the choice set. In a sense, our model provides a bridge between the two dominating structural labour supply models in the literature, the early generation “continuous” Hausman labour supply model and the new discrete choice approach based on random utility functions.

In the basic linear Hausman model, the individuals choose bunches of consumption and hours worked which maximise their utilities characterised by strictly positive marginal utilities with respect to consumption and leisure given their budget constraints. The constrained utility maximisation implies a labour supply function, which depends on the net market wage, the net non-labour income and individual preferences. In this framework, income taxes affect negatively labour supply via the substitution effect (taxes decrease the net market wage, which increases the opportunity cost of leisure) and positively via the income effect (a decrease in net income tend to cause individuals



to supply more labour in order to partially compensate for the lost income). Participation in paid work follows from a simple corner solution.

Although still used, the early generation Hausman model has been relentlessly criticized for several reasons. First, the traditional Hausman model relies on the assumption of (piecewise) linearity and convexity of the budget sets. This assumption is clearly violated if certain transfers get lost immediately on taking up a job, which is usually the case. As a consequence, the wage earned during the first few hours worked does not compensate for the discrete loss in benefits and thus the reservation wage is undefined. This limitation applies equally to more general, non-parametric estimation techniques (Blomquist and Newey (2002)). Fixed costs of taking up a job can also be introduced, yet this additional source of non-convexity poses a new challenge that is difficult to resolve (see e.g. Bourguignon and Magnac (1990)). van Soest and Das (2001) argue that non-convexities generally imply too restrictive and implausible forms for preferences.

Second, the Hausman model implicitly assumes quasi-concave utility function, which is typically rejected by empirical studies (see e.g. Blundell and MaCurdy (1999)). As argued by MaCurdy et al. (1990), the Hausman model implicitly imposes restrictions that generate a positive Slutsky effect at all internal points of the budget constraint. The coherency of the model thus implicitly limits the range of elasticities that can be obtained.

Third, estimates derived from the Hausman model exhibit a notable heterogeneity across studies and often yield unreliable elasticity measures at both the extensive and the intensive margins. For married men, there are a number of studies in which the income-compensated (Hicksian) wage elasticity of hours of work is estimated to be negative (see Pencavel (1987) for a survey of results for men). In contrast, the uncompensated wage elasticities of female labour supply are, in many cases, unrealistically high, often much larger than one (see Killingsworth and Heckman (1987) for a survey for women).<sup>1</sup> Although the direct comparison of the results across studies is often difficult because most of them focus on special subgroups and due to methodological differences – e.g. there is a substantial heterogeneity in the way after-tax wages are controlled for, if at all –, the meta-analysis of Chetty et al. (2013) provides a “new consensus estimate” of Hicksian extensive margin elasticities of 0.25. At the intensive margin, however, the Hausman model proved to be unable to accurately predict the hours worked distribution representative of actual hours as it does not take into account that very few observations exist with a small positive number of hours worked.

Finally, these studies usually focus on either taxes or transfers. As argued by Blundell (2012), it is important to take taxes and transfers into account simultaneously and combine them into effective

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<sup>1</sup>see also the more recent survey by Evers et al. (2008). The papers by Blundell and MaCurdy (1999), Blundell et al. (2007), Keane (2011) and Meghir and Phillips (2011) also survey the empirical literature based on the continuous approach, but also include some more recent evidence from discrete choice methodology.

tax wedges. Besides influencing non-labour income (income at zero hours worked), transfers also show characteristics resembling marginal and average (participation) tax rates. Suppose that a certain benefit is means tested with a gradual phaseout. For example, every extra income earned as wage reduces transfers by 20%. In that case, it is equivalent to a 20% extra marginal tax rate. Once the individual has lost all of this means tested benefit, lost transfers become similar to a participation tax rate: the total amount of lost transfers decreases the payoff from work, just like the participation tax rate does.

Since the seminal work of van Soest (1995), the discrete choice approach to modelling labour supply has become increasingly popular.<sup>2</sup> The discrete choice model is based on the concept of random utility maximization in which individuals choose a single option among a finite set of alternative hours worked. The model includes inactivity (zero hours worked) as one of the options so that both extensive and intensive margins are simultaneously estimated. In contrast to the Hausman approach, this method does not explicitly model labour supply. Instead, the estimation is carried out directly on the utility specification and labour supply is derived from individuals' or households' utility functions. In principle, the model can accommodate a very wide range of preference specifications with few restrictions. Most importantly, tangency conditions and the convexity of preferences need not be imposed.<sup>3</sup> The model only requires to impose the monotonicity in consumption and to accurately specify the consumption-leisure preference function. Moreover, the simultaneous presence of both taxes and transfers are relatively easily taken into account. Nevertheless, as argued by Dagsvik et al. (2014), the discrete choice model is similar to the standard textbook approach from a theoretical perspective as both are derived from the same theory of consumer behaviour. The only novelty is that the set of feasible hours of work is finite, and the state-dependence of transfers (which is in fact just another aspect of nonlinearity) can be taken into account.<sup>4</sup>

We base our modelling strategy on the mixture of the two previous approaches. We extend the standard labour supply model by incorporating both the marginal and participation tax rate aspect of transfers. We achieve this by building on the concept of the discrete choice models in which potential jobs are restricted to a discrete set of fixed hours worked. The labour supply choice of individuals is determined by comparing their utility level at each possible choice. However, we do not specify and estimate a given utility function as in random utility models. Instead, we derive and estimate individuals' labour supply function at the extensive margin.

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<sup>2</sup>However, it is generally less known that Zabalza et al. (1980) already used the discrete approach before van Soest (1995).

<sup>3</sup>As shown by Heim and Meyer (2004), it is senseless to impose the convexity of preferences as fixed costs of work cannot be disentangled from preference parameters.

<sup>4</sup>In parallel with structural labour supply models, there is also a growing literature using program evaluation methodology. Though such approaches are capable of precisely estimating the impact of a particular tax or transfer reform episode, they are not suitable for evaluating the impact of future (hypothetical) scenarios. See Moffitt (2002) for a review of this strand of the literature.

The goal of this paper is to provide a model for explaining individuals' labour supply at the extensive margin, i.e. the probability of choosing to work any positive amount of hours among the set of possible alternatives instead of working zero hours. In our model, the baseline category is necessarily inactivity and the utility derived from any other alternative is compared to the utility from staying out of the labour market. We start by restricting the choice set to zero hours worked and full time job. In this case, the individual will choose to work if his gains to work, defined as the difference between the net wage and the amount of transfers lost when working, is higher than his "reservation gains to work". The underlying theory - presented in Section 2.2 - leads to a binary dependent variable equation which relates participation probabilities to the gains to work from a full time job, the total amount of non-labour income (including the amount of transfers one gets or would get at zero hours worked) and other individual characteristics. The gains to work and the non-labour income are both influenced by taxes and (lost) transfers. Similarly to the standard discrete choice model, the analytical expression for the selection probabilities in the presence of several possible choice of hours worked is then derived using the axiom of independence of irrelevant alternatives (IIA) (see McFadden (1974)). Introduced by Luce (1959), the IIA states that the relative odds of one alternative being chosen over a second one is independent of the presence or absence of any other alternatives. Under this assumption, individuals' extensive labour supply decisions can be represented by a multinomial logit model. The strong IIA assumption can be relaxed by specifying a mixed logit model, which allows for correlation in unobserved factors over time and across alternatives.

To test our model, we carry out our estimations on the Hungarian Household Budget Survey (HKF). Our dataset contains detailed income and consumption measures of individuals for the years 1998-2008. This period featured numerous policy measures on both income tax rates and transfers, providing us with sufficient exogenous cross-sectional and time-series variation in non-labour income and gains to work. Hungary provides a particularly interesting case for estimating the determinants of labour market activity. Participation rates are among the lowest in the EU: the 2008 figure of 61.5% was the second lowest in the EU (after Malta), 9.2 percentage points lower than the EU28 average. By 2014 the gap narrowed to 6.8 percentage points, still yielding a position of 24/28.<sup>5</sup> This has been often identified as a key bottleneck to real convergence (see e.g. Kátay (2009)). On the other hand, the inconvenient of the dataset is that part-time work in Hungary is relatively rare: according to Eurostat, the share of part-time employees was less than 5% in our sample period. Therefore, we assume that there are only two labour-market states, active and inactive and only test the model with these two possible choices. Future research is needed to test the model with several alternatives.

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<sup>5</sup>The source is Eurostat, population between 15 and 64.

The estimation process - described in Section 2.3.1 - follows the often-used three step procedure, as e.g. in Kimmel and Kniesner (1998). The key element of the identification is the choice of labour demand shifters, i.e. the variables which have no (or negligible) impact on labour supply directly, but strongly impact the wage and hence impact activity indirectly. We argue that county dummies and experience are such variables. This latter is proxied by age, after controlling for its labour supply components by using a large set of lifecycle indicators.

Section 2.3.2 presents the estimation results. We find that a single equation can explain a large heterogeneity of individual responsiveness to taxes and transfers: there are marked differences in the conditional marginal effects among subgroups, driven partly by a composition effect (due to the nonlinearity of the probit function), and partly by a different share of lost transfers in the gains to work. The most responsive subgroups are low-skilled, (married) women at child-bearing age and elders,<sup>6</sup> while prime-age individuals with tertiary education are practically unresponsive to tax and transfer changes at the extensive margin. Our estimates imply an aggregate labour supply elasticity of 0.28, quite in line with the 0.25 “consensus estimate” of Chetty et al. (2013). As a quantitative illustration, we also fed the main changes of the Hungarian personal income tax system of 2012 into this framework. As a result, aggregate labour supply decreases by 0.97 percentage points. This is a sum of a 2.09 percentage points decline due to the elimination of the employee tax credit, a 0.34 percentage points decrease due to higher social contributions and an offsetting increase of 1.51 percentage points due to tax rate cuts.<sup>7</sup>

## 2.2. Theory

### 2.2.1. The continuous model of labour supply

The standard neoclassical theory of individual labour supply is based on the consumption-leisure trade-off with a budget constraint and a limited amount of time that the individual can allocate to work and leisure. Individual  $i$  chooses his labour supply that maximises his utility of the form:

$$U_i = U(c_i, 1 - l_i) \tag{2.1}$$

where  $c_i$  represents consumption and  $l_i$  is labour. The total time endowment is normalized to 1, so leisure is  $1 - l_i$ . The basic Hausman model simply derive the optimal labour supply given the utility function (2.1) and the budget constraint stating that consumption must equal total income:

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<sup>6</sup>As argued for example by Kátay (2009), Hungary’s labour participation deficit compared to other EU members is mostly due to these special groups.

<sup>7</sup>The sum of the effects of these measures may differ from the total effect due to interactions.

$c_i = w_i l_i + T_i$ , where  $w_i$  is net wage and  $T_i$  denotes non-labour income. The optimality condition can be written as:

$$U_{1-l}(c_i, 1 - l_i) = w_i U_c(c_i, 1 - l_i)$$

In this framework, the participation decision simply follows from the corner solution. The reservation wage is the lowest wage that induces nonzero hours worked. It corresponds to the case where  $1 - l_i = 1$ , i.e., tangency occurs exactly at zero hours worked. Then  $c_i = T_i$ , so

$$w_i^{res} = \frac{U_{1-l}(T_i, 1)}{U_c} \quad (2.2)$$

defines the reservation wage. The participation decision is then determined by  $w_i > w_i^{res}$ . Loglinearising the right hand side, or working with specific functional form for  $U$ ,<sup>8</sup> we get

$$\log w_i > \log \chi + \Psi \log T_i \quad (2.3)$$

To allow for individual heterogeneity, we assume that individuals differ in their utility of leisure term ( $\chi$ ) and then expand  $\log \chi$  as  $Z_i \Theta' + \varepsilon_i$ , where  $Z_i$  is a vector of observable individual characteristics and  $\varepsilon_i$  is an i.i.d. error term. It follows that the probability  $P_i$  of an individual  $i$  working given a wage offer  $w_i$ , non-labour income  $T_i$  and individual characteristics  $Z_i$  is

$$P_i = G(\gamma \log w_i + Z_i \theta' - \psi \log T_i) \quad (2.4)$$

with  $G$  being the cumulative distribution function (CDF) of the standard normal distribution if  $\varepsilon_i$  is assumed to follow a centred normal distribution or the CDF of the logistic distribution if the errors are distributed as logistic.

When adding taxes and transfers, two considerations have to be taken into account. On the one hand, one has to modify the wage rate by the effective tax rate, including taxes, social contributions, and the phaseout of social transfers (if applicable). On the other hand, there are certain transfers which get lost immediately at taking up any job. In such a case, there is a discrete downward jump  $\Delta T$  in  $T$  for any nonzero hours worked. One could try to redefine the reservation wage similarly as before, being the level that could still induce an epsilon amount of work. This is, however, not feasible: from Roy's identity, the welfare gain from a marginal wage increase is the same as the income gain from the extra income due to the higher wage. But there is no such income gain at zero

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<sup>8</sup>The standard separable CES form  $\frac{e^{1-\psi}-1}{1-\psi} + \chi \frac{(1-l)^{1-\phi}-1}{1-\phi}$  yields exactly (2.3). The growth-consistent utility function of the form  $\frac{(c \cdot \exp(g(1-l)))^{1-\psi}-1}{1-\psi}$  also yields (2.3), with an extra constraint of  $\psi = 1$ .

hours worked, so the income equivalent gain is zero, while there is a nonzero income loss due to the drop in  $T$ . In other words, the nonlinearity of the budget constraint at zero hours worked makes the reservation wage infinite (this can also be established formally by total differentiation).

### 2.2.2. The random utility model

The discrete choice approach provides a convenient way of solving the problem related to the nonlinear budget constraint. van Soest (1995) proposes to restrict the labour supply decision to a finite set of possible job sizes, including zero hours worked. This also implies that an interior solution at a too low number of hours is infeasible. Restricting the choice set considerably simplifies the model and provides a simple yet very general way of modelling individuals' or households' labour supply decision.

In this framework, agents' preferences are represented by a random utility model. The utility  $U_{ij}$  provided to the individual  $i$  by working  $l_j$  hours is composed of a deterministic component  $V_{ij}$ , which can be estimated using observed characteristics, and a stochastic error component  $\varepsilon_{ij}$  that is uncorrelated with the deterministic part:

$$U_{ij}(c_{ij}, 1 - l_j) = V(c_{ij}, 1 - l_j, Z_i, \Theta_j) + \varepsilon_{ij} \quad (2.5)$$

where  $Z_i$  is a set of individual characteristics and  $\Theta_j$  are preference parameters. Each individual chooses the alternative which provides the highest level of utility. The probability  $P_{ij}$  of individual  $i$  choosing the alternative  $j \in \{0, \dots, J\}$  is given by:

$$\begin{aligned} P_{ij} &= P[U_{ij} > U_{ij'}, \forall j' \neq j] \\ &= P[V(c_{ij}, 1 - l_j, Z_i, \Theta_j) + \varepsilon_{ij} > V(c_{ij'}, 1 - l_{j'}, Z_i, \Theta_{j'}) + \varepsilon_{ij'}, \forall j' \neq j] \end{aligned} \quad (2.6)$$

McFadden (1974) shows that if the error term follows an i.i.d extreme value type 1 distribution with c.d.f.  $\exp\{-\exp(-\varepsilon_{ij})\}$ , then the selection rule in eq. (2.6) implies that the probability of individual  $i$  choosing the alternative  $j$  can be represented by a multinomial logit specification:

$$P_{ij} = \frac{\exp\{V_{ij}\}}{\sum_{j'=0}^J \exp\{V_{ij'}\}} \quad (2.7)$$

The central assumption behind McFadden's multinomial logit is the strong axiom of independence of irrelevant alternatives (IIA), which states that the relative odds between two alternatives are the

same, no matter what other alternatives are available in the set of choices.<sup>9</sup> That is, the log-odds ratio between two alternatives  $j$  and  $j'$  is  $\ln(P_{ij}/P_{ij'}) = V_{ij} - V_{ij'}$ , which does not depend on any alternatives other than  $j$  and  $j'$ .

The parameters of the model can be estimated by maximum likelihood techniques. Utility being ordinal, it is not possible to recover all estimates of  $\Theta_j$ . A baseline category is selected (usually inactivity, category  $j = 0$ ) against which other alternatives are assessed. The model can thus be seen as consisting of  $J$  simultaneous logit equations of the form  $\ln(P_{ij}/P_{i0}) = V_{ij} - V_{i0} + \tilde{\varepsilon}_{ij}$ , where  $\tilde{\varepsilon}_{ij} = \varepsilon_{ij} - \varepsilon_{i0}$  follows a standard logistic distribution.<sup>10</sup>

The major weakness of the multinomial logit model is the highly restrictive IIA assumption. For example, the IIA assumption is violated as soon as an important variable is missing from the model. To see this, let's assume that a relevant individual specific variable  $z_i$  with true parameters  $\beta_j$  is unobserved. The error terms become  $\varepsilon_{ij} = \beta_j z_i + \varepsilon'_{it}$ , which are now necessarily correlated because of the common influence of the omitted variable. In practice, the cross-correlation of the error terms is highly probable. For instance, an individual who places a higher value on working full time is also likely to have a higher probability of accepting a part time job than an other, observationally equivalent individual with a strong preference for leisure.

There are several ways to overcome this limitation, the most widely used in the structural labour supply literature is the mixed logit estimation procedure (see Train (2003) for a detailed description). The approach get rid of the main limitations of standard multinomial logit by allowing for correlation in unobserved factors across alternatives and over time and consequently, it can accommodate very flexible substitution patterns and random taste variation. The mixed logit model allows a subset  $B$  of the parameters – that can include the alternative-specific constants or, in the extreme case, all the parameters of the model – to vary randomly across individuals. Conditional on knowing the the true values of  $B$ , the choice probabilities can be expressed as in (2.7). The unconditional mixed logit probabilities are the integrals of standard logit probabilities over a density of parameters:

$$P_{ij} = \int L_{ij}(B)f(B)dB \quad (2.8)$$

<sup>9</sup>The two other axioms are much less restrictive. First, the strict positivity constraint of the probabilities ( $P_{ij} > 0, \forall j$ ) is not restrictive as there is no meaningful difference between zero probability and tiny positive probability. Second, as argued by McFadden (1974), the axiom of additive separability ( $U(Z, k, b) = u(Z, k) - u(Z, b)$ ), a.k.a. the axiom of irrelevance of alternative set, directly follows from the first two axioms.

<sup>10</sup>Note that the difference of two extreme value type 1 distributed variables follows a standard logistic distribution.

where  $L_{ij}$  is the logit probability in (2.7) evaluated at different values of  $B$ , with the weights given by the density  $f(B)$ . Eq. (2.8) is a multi-dimensional integral that has no closed-form solution. The model can be estimated using maximum simulated likelihood technique (see Train (2003)).

### 2.2.3. The gains to work model

The gains to work model of labour supply that we propose in this paper is based on similar assumptions to the traditional discrete choice approach. Specifically, we assume that the labour supply decision is constrained to a set of fixed job sizes. However, instead of explicitly estimating the structural parameters of the underlying utility function, we derive a labour supply function in which the individual compares the financial gains from accepting a job (the gains to work) with a certain threshold that can be interpreted as the “reservation gains to work”.

Similarly to the random utility model, the selection rule is given by equation (2.6). For each possible choice, the consumer is bound by the following budget constraint:

$$c_{ij} = w_i l_j + T_i - \Delta T_{ij} = W_{ij} + T_i, \forall j \quad (2.9)$$

where  $w_i$  is the net hourly market wage offered to individual  $i$ ;  $\Delta T_{ij}$  is the total amount of all the hypothetical social transfers the individual gets (or would get) at zero hours worked and loses when working  $l_j$  hours; and  $T_i$  is the total net non-labour income, including social transfers ( $\Delta T_{ij}$ ) and other non-labour income such as, for example, capital income or income of the other members of the household. We define  $W_{ij} = w_i l_j - \Delta T_{ij}$  as the financial gains from accepting to work  $l_j$  hours instead of being inactive.

Let's assume, first, that individuals' decision problem is reduced to choosing between working full time ( $j = 1$ ) or being inactive ( $j = 0$ ). The reservation wage is thus set by the following comparison:

- Do not work: then  $c_{i0} = T_i, 1 - l_0 = 1$  and welfare is  $U_{i0}(T_i, 1)$ .
- Work  $l_1$  hours: then  $c_{i1} = w_i l_1 + T_i - \Delta T_{i1} = W_{i1} + T_i$  and welfare is  $U_{i1}(W_{i1} + T_i, 1 - l_1)$

The selection rule in (2.6) implies that the individual will choose to participate in the labour market if and only if:

$$U_{i1}(W_{i1} + T_i, 1 - l_1) > U_{i0}(T_i, 1) \quad (2.10)$$

To derive a formal expression for the probability of being active, we linearise the left hand side of (2.10):



$$U_{i1}(W_{i1} + T_i, 1 - l_1) \approx U_{i1}(T_i, 1 - l_1) + W_{i1}U'_{i1}(T_i, 1 - l_1)$$

so the comparison becomes:

$$W_{i1} > \frac{U_{i0}(T_i, 1) - U_{i1}(T_i, 1 - l_1)}{U'_{i1}(T_i, 1 - l_1)} \quad (2.11)$$

Since  $U_{i1}(T_i, 1 - l_1) < U_{i0}(T_i, 1)$  and  $U'_{i1}(T_i, 1 - l_1) > 0$ , the right-hand side of (2.11) is positive. It represents the "reservation gains to work", i.e. the minimal difference between the net wage and the amount of lost transfers required by the individual to accept a job offer. Equation (2.11) is analogous to the reservation wage condition in (2.2) and can be seen as the discretised version of the Hausman model.

Loglinearising the right hand side, or working with the same specific functional forms for  $U$  as before, the individual works if

$$\log W_i > \log \chi - \psi \log T_i$$

Expanding  $\log \chi$  as  $Z_i\Theta' + \varepsilon_i$  yields a binary dependent variable model:

$$P_{i1} = G(\gamma \log W_i + Z_i\theta' - \psi \log T_i). \quad (2.12)$$

Equation (2.12) is similar to the final equation (2.4) of the Hausman approach with two important differences. First,  $W_i$  in (2.12) represents the gains to work (from a full time job), as opposed to the net wage  $w_i$ . Second,  $T_i$  includes the hypothetical amount of transfers one gets (or would get) at zero hours worked. When compared with random utility models, we estimate a labour supply function by approximating the "reservation gains to work" by a set of observed individual characteristics instead of explicitly parametrising individuals' utility functions.

Let's turn now to the case with several possible alternative choices. In the same vein as in McFadden's model, we make use of the IIA axiom (Luce (1959)) stating that the choice between two alternatives is independent of the presence or absence of any other alternative. In other words, the relative log-odds between any of the alternatives and the base outcome (inactivity,  $j = 0$ ) does not depend on the attributes of any other alternative. Therefore, assuming that the error terms  $\tilde{\varepsilon}_{ij}$  follow a standard logistic distribution, the choice probabilities can be modelled as  $J$  independent

logit equations of the form:

$$\ln(P_{ij}/P_{i0}) = \gamma_j \log W_{ij} + Z_i \theta'_j - \psi_j \log T_i + \tilde{\varepsilon}_{ij}, \forall j$$

which yields McFadden's multinomial logit specification:

$$\begin{aligned} \sum_{j=1}^J P_{ij} + P_{i0} &= \left( \sum_{j=1}^J \exp \{ \gamma_j \log W_{ij} + Z_i \theta'_j - \psi_j \log T_i \} + 1 \right) P_{i0} = 1 \\ \Rightarrow P_{i0} &= \frac{1}{1 + \sum_{j=1}^J \exp \{ \gamma_j \log W_{ij} + Z_i \theta'_j - \psi_j \log T_i \}} \\ \Rightarrow P_{ij} &= \frac{\exp \{ \gamma_j \log W_{ij} + Z_i \theta'_j - \psi_j \log T_i \}}{1 + \sum_{j'=1}^J \exp \{ \gamma_{j'} \log W_{ij'} + Z_i \theta'_{j'} - \psi_{j'} \log T_i \}} \end{aligned} \quad (2.13)$$

Imposing the IIA property provides a convenient closed form solution to the choice probabilities. However, this assumption is not more (and not less) credible than in random utility models. Indeed, unobserved differences in preferences may affect individuals' reservation gains to work for each alternative choice, which may create correlation across alternatives. To get rid of the IIA property, the same techniques can be used as in random utility models. For instance, a mixed logit model can be estimated using maximum simulated likelihood technique (Train (2003)) in which the conditional logit probabilities ( $L_{ij}$ ) are given by (2.13) and the unconditional mixed logit probabilities are given by (2.8).

Practical and econometric issues of estimating the gains to work model are the same as those of the random utility model. First,  $\Delta T_{ij}$  and, consequently,  $T_i$  and  $\log W_{ij}$  are not directly observable for each alternative even if the hourly market wages are known, since we do not know the amount of the transfer an individual would be entitled to if he increased or decreased his working hours. Using individual characteristics and the welfare system's details (for every given year), however, one can back up  $T_i$  and  $\Delta T_{ij}$ . Just like in the random utility model, this step essentially requires the detailed coding of the full transfer system, basically a microsimulation model.<sup>11</sup> For those who do not work,  $T_i$  is observed and  $\Delta T_{ij}$  are determined based on their characteristics and welfare regulations for the given year. For those who work,  $T_i$  is calculated by assuming that the individual has no income from work and by applying welfare rules, while  $\Delta T_{ij}$  is the simulated loss in transfers if the individual works  $l_j$  hours.

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<sup>11</sup>In the random utility model, we need to estimate the total disposable income (the proxy for  $c_{ij}$ ) for each alternative choice.

Second, the wages for non-employed have to be estimated. Usually a Heckman method is used, however, there is a considerable heterogeneity in the way that wages are treated in discrete choice models. The two methods frequently used in the empirical literature is predicting wages for the full sample and using predicted wages only for non-workers. Both methods have advantages and disadvantages – see e.g. MaCurdy et al. (1990) for a discussion and Löffler et al. (2014) for the comparison of the implied elasticities using one method or the other –, and this paper does not seek to add to this polarised discussion. This issue is not related to the modelling strategy, it is equally relevant in both the random utility model and the gains to work model of labour supply. In the empirical part of the paper, we opted for the former method, i.e. using predicted wages – or, more precisely, predicted gains to work – for all individuals because it can at least partly resolve the endogeneity problem between wage rates and labour supply decisions, provided that the excluded instruments are valid. We argue that county dummies and – after controlling for its labour supply components by using a large set of lifecycle indicators – work experience are such variables.

### 2.3. Labour supply at the extensive margin in Hungary

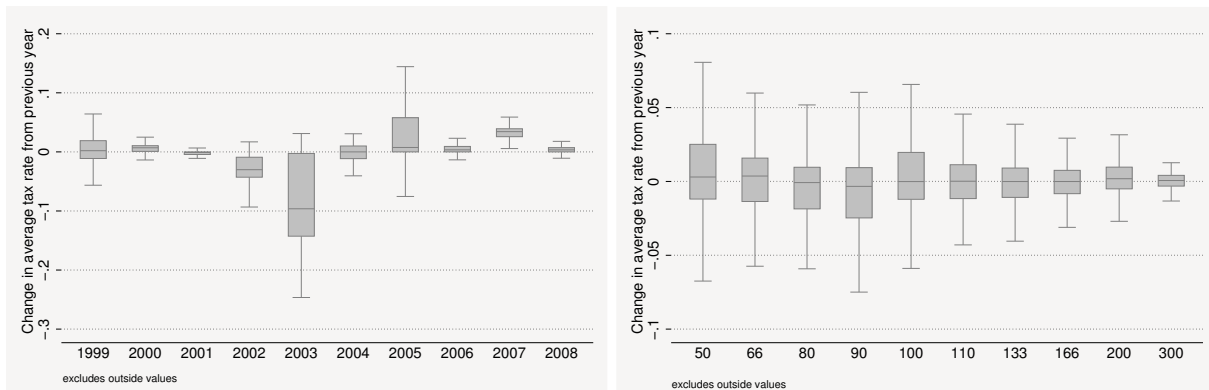
We test our model on Hungarian Household Budget Survey (HKF). The dataset contains detailed income and consumption measures of broadly 25,000 individuals per year in a repeated cross-section between 1998 and 2008.<sup>12</sup> We detail the major tax expenditure and cash transfer items in the Appendix. With one exception, the database contains all the relevant information to deduct the counterfactual transfer entitlements or losses of each individual. The exception is the work history of individuals, on which certain transfers depend (for example, eligibility to the more generous maternity support schedule GYED). To resolve this issue, we used a predicted value based on the Labour Force Survey database (a conditional expectation based on observable characteristics).

The sample period features numerous policy measures on both income tax rates and transfers, providing us with sufficient exogenous cross-sectional and time-series variation in non-labour income and gains to work. Starting with taxes, Figure 2.1 shows how individuals' average tax rates would have changed if their real income had remained unchanged over time. For each individual and for each year, we simulated the change in the average tax rate (ATR) that the taxpayer would have faced if his real income had not changed. Panel A displays the lowest, the highest and the mean ATR change by year, as well as the lower and the upper quartile of the distribution. It is apparent that there were some income tax changes every year, while major changes occurred in 1999 and between 2002 and 2005. Panel B depicts the main characteristics of the distribution of ATR changes

<sup>12</sup>Though the dataset has some potential for following individuals in time, it is very difficult to make the actual connections between consecutive waves and then make the panel dataset representative.

by income categories. The graph reveals that tax changes during the period between 1998 and 2008 affected low income earners to a greater extent than high income earners.

**Figure 2.1.:** Variation in the changes in average tax rates (ATR)\*



(a) Yearly changes in ATR

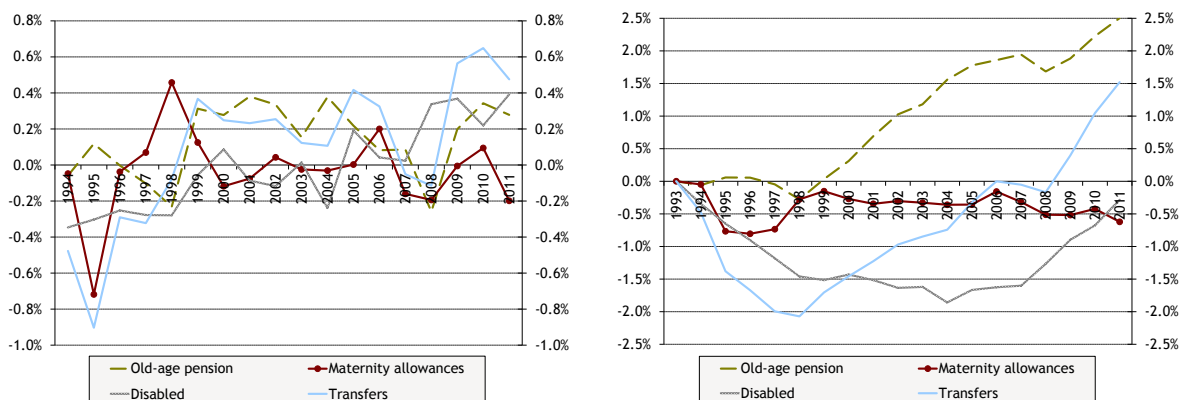
(b) Changes in ATR by income categories (% of average income)

\* Source: Household Budget Survey, own calculations

Graphs show the yearly changes in average tax rates between 1998 and 2008 for the individuals observed in 2008, assuming that their real income did not change during this period. Each column shows the lowest, highest and the mean ATR change, and the upper and lower quartile of the distribution.

As for transfers, Figure 2.2 – adopted from Kátay and Nobilis (2009) – illustrates the impact of various transfer reforms on the Hungarian participation rate. The authors decomposed the changes in the aggregate labour force participation rate over time into changes in the labour force participation behaviour of different population groups and changes in each group's population share. Their results reveal that the major welfare reforms in Hungary have had a significant impact on the share of benefit recipients in the population, implying that transfer changes do impact the participation rate. As shown in Figure 2.2, the temporary tightening of the childcare benefit system between 1996 and 2000, the gradual increase of old-age retirement age since 1998 and the tightening of the conditions of the disability pension since 2007 have all led to significant changes in the aggregate participation rate.

Part-time work in Hungary is relatively rare: according to Eurostat data, the share of part-time employees was less than 5% in our sample period. Consequently, we assume that there are only two labour-market states, active and inactive, and we only test the model with these two possible choice. Future research is needed to test the model with several possible alternative hours worked.

**Figure 2.2.:** Decomposition of the aggregate participation rate\***(a)** Year-to-year percentage point changes**(b)** Cumulative changes

\* Source: Kátay and Nobilis (2009), updated

### 2.3.1. Econometric issues

We estimate equation (2.12) using a three-step procedure, where the first two steps are a Heckman model aiming at predicting wages (or gains to work) for all individuals or only for non-workers, as discussed in the previous section. Individuals' wages are regressed on a vector  $X_i$  containing relevant individual characteristics. Note that the two vectors  $X_i$  and  $Z_i$  may overlap, but there can be elements in each of them which are excluded from the other set. This regression, however, is run on a nonrandom sample, since the employment and the wage error terms might be correlated. The solution is thus to adopt a Heckman-type correction, yielding a three-step procedure.<sup>13</sup> Given that our wage variable is gains to work, one can follow two alternatives. In variant A, the procedure is the following:

1. Run a reduced form probit

$$P_i(\text{employed}) = \Phi(X_i \beta'_{RF} + Z_i \theta'_{RF} - \psi_{RF} \log T_i)$$

2. Use the inverse Mills ratio  $\lambda(x) = \frac{\phi(x)}{\Phi(x)}$  as a correction in the log gains to work regression:

$$\log W_i = X_i \beta' + \delta \lambda(X_i \hat{\beta}'_{RF} + Z_i \hat{\theta}'_{RF} - \hat{\psi}_{RF} \log T_i) + \mu_i$$

<sup>13</sup>Kimmel and Kniesner (1998) follows such an approach, for example.

3. Use the predicted  $\widehat{\log W}_i = X_i \widehat{\beta}'$  in the structural probit equation:

$$P_i(\text{employed/active}) = \Phi(\gamma \widehat{\log W}_i + Z_i \theta' - \psi \log T_i) \quad (2.14)$$

Notice that here  $X \supseteq Z$ , since there is practically no observable characteristics which would not be related to transfer measures and thus to  $\log W$ .

In variant B, we slightly modify the previous procedure:

1. Run a reduced form probit

$$P_i(\text{employed}) = \Phi(X_i \beta'_{RF} + Z_i \theta'_{RF} - \psi_{RF} \log T_i)$$

2. Use the inverse Mills ratio  $\lambda_i(x) = \frac{\phi(x)}{\Phi(x)}$  as a correction in the wage (more precisely: yearly income) regression:

$$\log w_i = X_i \beta' + \delta \lambda(X_i \widehat{\beta}'_{RF} + Z_i \widehat{\theta}'_{RF} - \widehat{\psi}_{RF} \log T_i) + \mu_i$$

3. If  $W_i$  is also lognormal with some mean and a variance  $\sigma_W^2$ , then one can show that  $E(\log W_i | X_i, Z_i) = \log(E(W_i | X_i, Z_i)) - \frac{1}{2} \sigma_W^2 = \log(e^{X_i \beta_1 + \frac{1}{2} \sigma_1^2} - \Delta T_i) - \kappa$ . Thus we can use the predicted  $\widehat{\log w}_i = X_i \widehat{\beta}'$ , add the standard error correction for lognormals, exponentiate, subtract  $\Delta T_i$  and take logs again to obtain the predicted  $\widehat{\log W}_i$  for the structural probit equation as in (2.14).

Four remarks are in order. The first is regarding endogeneity and measurement error of the gains to work variable. In the structural probit,  $\log W$  can be endogenous, since the wage error term can be correlated with the participation decision error term. Moreover,  $\log W$  can also contain measurement error: in case of an individual working only for some part of the year, his reported wage is less than the true annual wage. Alternatively, unreported wage income can also lead to a mismeasurement of wages. Notice, however, that using predicted wages or gains to work for all individuals, we are in fact running an IV-probit. This strategy offers a remedy to both of these problems, provided that there are variables in  $X_i$  which are excluded from  $Z_i$ .<sup>14</sup>

The second issue is whether the selection correction is identified only through a functional form assumption. This is indeed the case when  $X \supseteq Z$  in the wage equation, since the inverse Mills ratio is then just a nonlinear reshuffling of the right hand side variables in the wage equation (variant A). On the other hand, the inverse Mills ratio does contain additional variation if  $X \not\supseteq Z$ , which is

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<sup>14</sup>Note that many empirical papers simply ignore the endogeneity of wages. See e.g. Löffler et al. (2014) for an overview.

the case in Variant B. This means that we are free from the functional-form criticism in Variant B, but it applies for the wage equation in Variant A. In that case, however, there is no alternative: if a variable impacts the participation equation directly, it is also likely to impact  $\log W$  at least through the change in transfers term  $\Delta T$ . For the structural probit equation (2.14) however, we are again on safe grounds: though the predicted  $\log W$  contains the variables  $X$ ,  $Z$  and their nonlinear combinations (in the inverse Mills ratio),  $X$  is excluded from the structural equation, so we are identifying  $\gamma$  from variations both in  $X$  and the inverse Mills ratio. In short, the key element of the identification method is the existence of controls for labour demand included in  $X_i$  and excluded from  $Z_i$ .

Third, the use of generated regressors in the third stage calls for an adjustment of standard errors. Usual Heckman correction implementations do incorporate necessary corrections for the second but not for the third step. In practice, such a correction often leads to minor changes; hence it is common to ignore the issue (Kimmel and Kniesner (1998) also follow this route). As one alternative, one could implement a full-blown correction of the third step standard errors, along the lines of Fernandez et al. (2001). We instead opted for bootstrapping the standard errors, which should be more robust in case of noisy data or misspecification problems.<sup>15</sup>

Finally, there is a tradeoff between adopting Variant A or B. The latter would seem more appealing, since it allows for  $X \not\subseteq Z$ , hence even the wage equation is free from functional form criticisms. The drawback, however, is that nothing guarantees that our estimated  $\hat{W}_i = e^{X_i \hat{\beta}_1 + \frac{1}{2} \hat{\sigma}_1^2} - \Delta T_i$  is positive, causing a nonrandom sample selection issue in our third step. One could produce better second stage regressions for  $\log w_i$ , taking for example the impact of the minimum wage into account.<sup>16</sup> That would mean, however, a Tobit-type truncated regression in the second stage, making our procedure even more complicated and potentially four-step. For this reason, we proceed only with Variant A; also recalling that although the wage equation is subject to a functional form criticism, it is much less of an issue in the structural probit equation.

Since our “wage” measure in the structural estimation is the gains to work, the calculation of regular wage elasticities requires one more step. The structural probit gives us a  $\log W$  coefficient  $\gamma$ . Since the probit is a nonlinear function, one has to evaluate it at a certain vector  $Z$  and  $\log T$  to obtain the marginal impact of a percentage change in the gains to work. This already implies some heterogeneity of elasticities. Even then, however, it is still the impact of a change in  $W$ , not  $w$ .

<sup>15</sup>In particular, our reported standard errors are calculated as the standard deviation of the point estimates from the three-step estimation procedure performed on 200 bootstrapped random samples (with replacement, and of the same size as the estimation sample).

<sup>16</sup>It was indeed the case in our sample that the predicted wage was too low for the low-skilled, where the minimum wage is often binding, and this made their predicted  $\hat{W}_i$  negative.

To obtain the impact of the wage itself, note that

$$\frac{\partial \log(w - \Delta T)}{\partial \log w} = \frac{\partial \log(e^{\log w} - \Delta T)}{\partial \log w} = \frac{e^{\log w}}{e^{\log w} - \Delta T} = \frac{w}{w - \Delta T}$$

so

$$\frac{\partial \Phi}{\partial \log w} = \frac{\partial \Phi}{\partial \log W} \frac{\partial \log W}{\partial \log w} = \frac{\partial \Phi}{\partial \log W} \frac{w}{w - \Delta T}. \quad (2.15)$$

Notice that the marginal effect of  $\log W$  gets magnified if  $w - \Delta T \ll w$ ; which is the case for transfer-dependent people (low skill, around retirement, etc.). This further increases the population heterogeneity of (net) wage elasticities.

The main left hand side variable is labour force participation,<sup>17</sup> though we also ran the same estimations with employment. All wage variables ( $w$  and  $W$ ) are based on annual net wage income, calculated from the gross wages reported by survey participants. The right hand side measures form two major groups: labour-supply shifters ( $Z_i$ ) and wage equation controls ( $X_i \setminus Z_i$ ). Following MaCurdy (1985), MaCurdy (1987) and Kimmel and Kniesner (1998), labour-supply shifters contain personal and family characteristics, while the vector  $X_i$  includes variables which determine the market wage (labour demand shifters). In particular, the first group consists of the following variables: log of non-labour income, education dummies, household head, mother with infant (< 3 years old), attending full-time education, household size (number of persons), pensioner, family status (husband, wife, child, single, divorced,...), age-group dummies (15-24, 25-49, 50-) and year dummies. The second group ( $X_i \setminus Z_i$ ) contains county dummies, and interactions of age and age square with education.

One needs to justify the choice for variables in  $X_i \setminus Z_i$ , since those variables serve both as instruments for treating endogeneity and measurement error issues about our wage measure, and also as a source of additional variation to identify the parameter  $\gamma$ . County dummies represent regional differences in economic conditions, which has an indirect effect on activity (through different wages) but no direct effect: two individuals with identical individual characteristics and wage but living in different regions should exhibit the same attitude towards economic activity. For the interaction of age and age square with education, our argument is the following. Age has two main effects on the likelihood of activity: one is through an impact on the lifecycle position (student, prime age and nearing retirement), and another through increased experience (an upward sloping relationship between age and wages). The first effect is a labour-supply shifter, which we capture by a large set of dummies that controls for individuals' lifecycle position, such as age-group, family status

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<sup>17</sup>It is the "most typical" status for the given year, self-reported by survey respondents. Unemployment is defined along the ILO classification.



(single, married, divorced...), attending full-time education, mother with infant and others. On top of that, we argue that an extra year has a negligible impact on labour supply directly, but it strongly impacts the wage and hence impacts activity indirectly (a labour demand shifter).

## 2.3.2. Results

This section reports and discusses our empirical results. We are mainly interested in how the gains to work and non-labour income affect labour force participation. With employment being the left hand side variable, we only report the results of the main specification but not the detailed conditional marginal effects by subgroups (they are available upon request). The main parameters of interest are the coefficient of gains to work and non-labour income (always in logs).

### 2.3.2.1. Overall participation elasticities

Table 2.1 displays our baseline results, following the econometric methodology of Variant A. Panel A reports the estimates for the structural probit equation (2.14). Most point estimates have the expected sign and are significant. A higher gains to work increases the probability of being active, while non-labour income has the opposite effect (both are in logs). From the additional controls (unreported but available upon request), education has a mixed but insignificant effect. Being a household head or having a larger family increases the probability of being active, while being a mother with small children, full-time student or pensioner decreases it. Age has the usual hump-shaped effect on activity. The results are quite similar when the left hand side variable is employment.

Since the probit function is nonlinear, the point estimates in Panel A are not indicative about the conditional marginal effect of variables of interest on activity. Panel B displays these numbers, evaluated at the sample means. Numbers here are already semi-elasticities: a 10% increase in the gains to work leads to a 2.9 percentage points increase in the probability of being active. As explained by equation (2.15), the same increase in the net wage (as opposed to the net wage minus transfers) leads to a potentially larger effect. The difference is quite substantial at the sample mean, as the effect is about 36% higher. The opposite happens with non-labour income: transfers are only part of them, so a 10% change in transfers implies a smaller increase in non-labour income.

**Table 2.1.:** Main results

|                   | (A) Estimation results           |                  |              |                  |
|-------------------|----------------------------------|------------------|--------------|------------------|
|                   | participation                    |                  | employment   |                  |
|                   | (1)                              | (2)              | (1)          | (2)              |
|                   | <i>coef.</i>                     | <i>std. err.</i> | <i>coef.</i> | <i>std. err.</i> |
| gains to work     | 0.820                            | 0.099            | 0.761        | 0.089            |
| non-labour income | -0.844                           | 0.110            | -0.702       | 0.098            |
|                   | (B) Conditional marginal effects |                  |              |                  |
|                   | <i>dy/dx</i>                     | <i>std. err.</i> | <i>dy/dx</i> | <i>std. err.</i> |
| gains to work     | 0.290                            | 0.028            | 0.301        | 0.031            |
| non-labour income | -0.298                           | 0.030            | -0.277       | 0.035            |
| net wage          | 0.395                            | 0.038            | 0.410        | 0.042            |
| transfer          | -0.136                           | 0.013            | -0.137       | 0.015            |

*Source: Household Budget Survey database, 1998-2008. Notes: Three-step estimates, as described in the paper. Standard errors are bootstrapped with 200 replications. Structural probit equation includes: log of gains to work, log of non-labour income, mother with infant (less than three years-old), full time student, education dummies (less than elementary school, elementary school, vocational, secondary education, tertiary education), age-group dummies (15-24, 25-49, >=50), pensioner, gender, head of household dummy, household size, family status dummies (single, married living together, married living separately, widow(er), divorced), household membership status dummies (husband, wife, companion, single parent, child, ascendant, other relation, non-relation, single), year dummies. Controls included in the reduced-form probit and the wage equation which are missing from the structural probit are: county dummies, interaction of age and age square with education dummies.*

The conditional marginal effects presented in Table 2.1 are not directly comparable to the ‘consensus’ 0.25 value of aggregate net wage elasticity reported by Chetty et al. (2013): these marginal effects indicate the effect of one percent increase in net wage on the “average individual’s” probability of being active (or on the participation rate) in percentage points, as opposed to the elasticity measures in Chetty et al. (2013) indicating the percentage change in total employment to the same shock. To produce the equivalent of the exercise by Chetty et al. (2013), one needs to increase the net wage of all individuals by one percent, calculate the change in their participation probability and then aggregate over the sample. To obtain the proper change in aggregate participation, prob-

ability changes are weighted by the appropriate sample weights. The resulting 0.28% increase in total employment implies an elasticity of 0.28, quite in line with the consensus.

### 2.3.2.2. Participation elasticities by subgroups

Next we look at the conditional marginal effects by subgroups to see how much they differ from each other. Table 2.2 presents two variants, a full and a restricted sample estimate. The full sample means that all observations are included (as in Table 2.1), but the marginal effects are evaluated at a subgroup-specific mean. The restricted sample means that the entire estimation procedure is carried out only on the subsample at hand, so even the structural probit estimates can be different.

**Table 2.2.:** Probit estimates and conditional marginal effects by subgroups

|                              |                            | full sample  |                  | restricted sample |                  |
|------------------------------|----------------------------|--------------|------------------|-------------------|------------------|
|                              |                            | (1)          |                  | (2)               |                  |
|                              |                            | <i>dy/dx</i> | <i>std. err.</i> | <i>dy/dx</i>      | <i>std. err.</i> |
| elementary<br>school or less | gains to work (probit)     | 0.820        | 0.099            | 0.583             | 0.082            |
|                              | non-labour income (probit) | -0.844       | 0.110            | -0.639            | 0.111            |
|                              | gains to work              | 0.212        | 0.064            | 0.175             | 0.085            |
|                              | non-labour income          | -0.218       | 0.068            | -0.192            | 0.101            |
|                              | net wage                   | 0.294        | 0.089            | 0.275             | 0.133            |
|                              | transfer                   | -0.093       | 0.028            | -0.109            | 0.053            |
| secondary<br>education       | gains to work (probit)     | 0.820        | 0.099            | 0.710             | 0.151            |
|                              | non-labour income (probit) | -0.844       | 0.110            | -0.715            | 0.165            |
|                              | gains to work              | 0.219        | 0.022            | 0.213             | 0.031            |
|                              | non-labour income          | -0.225       | 0.024            | -0.214            | 0.034            |
|                              | net wage                   | 0.310        | 0.031            | 0.286             | 0.041            |
|                              | transfer                   | -0.118       | 0.012            | -0.098            | 0.014            |
| tertiary<br>education        | gains to work (probit)     | 0.820        | 0.099            | 0.915             | 0.323            |
|                              | non-labour income (probit) | -0.844       | 0.110            | -0.856            | 0.326            |
|                              | gains to work              | 0.110        | 0.012            | 0.130             | 0.029            |
|                              | non-labour income          | -0.113       | 0.012            | -0.121            | 0.031            |
|                              | net wage                   | 0.139        | 0.015            | 0.156             | 0.035            |
|                              | transfer                   | -0.045       | 0.005            | -0.043            | 0.010            |

Notes: Column (1) reports probit estimates and conditional marginal effects computed from the estimation on the full sample and evaluated at the subgroup-specific mean values of the covariates. Column (2) reports similar marginal effects, but computed from the estimations on the restricted samples.

Notice that the net wage (or even the gains to work) elasticity of activity is lower for tertiary educated individuals than for people with primary or secondary education even in the full sample estimation case, when the only reason is a different conditional mean of the subgroups. The probit estimates somewhat differ between the full and the restricted sample, though the latter is often less precisely estimated. Still, the conditional marginal effects are quite similar. This result is noteworthy, as it means that one can explain the heterogeneity of participation elasticities without an underlying difference in the utility functions (i.e., the parameters  $\gamma$  and  $\psi$  in equation (2.14)).

Table 2.3 further explores the prime-age sample, checking whether education status also matters there. The low overall elasticity of this age group splits into a sizeable elasticity for the "elementary school or less" group (a group which is also highly welfare dependent) and a smaller but still significant number for prime-age individuals with secondary education. Estimations suggest that prime-age higher educated individuals are inelastic to tax and transfer changes at the extensive margin. The restricted samples yield similar though smaller differences, both for structural probit parameters and conditional marginal effects.

Table 2.4 displays the conditional marginal effects for the two remaining main welfare dependent social groups, the elderly and women of child-bearing age. The group of age above 50 exhibits a very substantial elasticity – this partly explains the large gap between the elasticity of the entire population and the prime-age group. This finding is quite important, as it shows that taxes and transfers have a strong impact on activity around retirement age, and that the tax and social insurance system can contribute to the large activity gap of the elderly in Hungary. Women at child-bearing age show a smaller wage elasticity, though they are still more responsive than the overall prime-age group. This is also true about the impact of transfers.

Finally, Table 2.4 also report results for the usual classification by sex and marital status. Consistently with most of the previous empirical findings, women are, in general, more responsive to tax and transfer changes than men. Married women, the group mostly studied in the literature exhibits the highest marginal elasticity, while married men seem to be the less responsive group.

In summary, we have found that wages, taxes and transfers have a large impact on the participation decision, particularly for elders, the low-skilled, married women and women at child-bearing age. Moreover, these differences can be largely explained by different group characteristics, leading to different conditional marginal effects of the same structural probit estimates, and also to a different multiplication of a net wage change into the change in the gains to work. Overall, we find a similar degree of heterogeneity in the participation elasticity to the one utilized in the calibration of Immervoll et al. (2007), ranging from 0.4 to zero.

**Table 2.3.:** Probit estimates and conditional marginal effects by subgroups, prime-age subsample

|  |                            | full sample  |                  | restricted sample |                  |
|--|----------------------------|--------------|------------------|-------------------|------------------|
|  |                            | (1)          |                  | (2)               |                  |
|  |                            | <i>dy/dx</i> | <i>std. err.</i> | <i>dy/dx</i>      | <i>std. err.</i> |
| full prime-age<br>sample                   | gains to work (probit)     | 0.820        | 0.099            | 0.646             | 0.122            |
|  | non-labour income (probit) | -0.844       | 0.110            | -0.620            | 0.129            |
|  | gains to work              | 0.088        | 0.010            | 0.086             | 0.008            |
|  | non-labour income          | -0.091       | 0.010            | -0.083            | 0.008            |
|  | net wage                   | 0.127        | 0.014            | 0.124             | 0.011            |
|  | transfer                   | -0.054       | 0.006            | -0.051            | 0.005            |
| prime-age,<br>elementary<br>school or less | gains to work (probit)     | 0.820        | 0.099            | 0.323             | 0.164            |
|  | non-labour income (probit) | -0.844       | 0.110            | -0.299            | 0.185            |
|  | gains to work              | 0.249        | 0.025            | 0.109             | 0.051            |
|  | non-labour income          | -0.256       | 0.026            | -0.101            | 0.058            |
|  | net wage                   | 0.409        | 0.040            | 0.180             | 0.085            |
|  | transfer                   | -0.194       | 0.019            | -0.084            | 0.041            |
| prime-age,<br>secondary<br>education       | gains to work (probit)     | 0.820        | 0.099            | 0.403             | 0.182            |
|  | non-labour income (probit) | -0.844       | 0.110            | -0.364            | 0.192            |
|  | gains to work              | 0.081        | 0.008            | 0.057             | 0.017            |
|  | non-labour income          | -0.084       | 0.008            | -0.051            | 0.019            |
|  | net wage                   | 0.122        | 0.012            | 0.084             | 0.025            |
|  | transfer                   | -0.054       | 0.005            | -0.036            | 0.011            |
| prime-age,<br>tertiary<br>education        | gains to work (probit)     | 0.820        | 0.099            | -0.206            | 0.420            |
|  | non-labour income (probit) | -0.844       | 0.110            | 0.217             | 0.400            |
|  | gains to work              | 0.038        | 0.003            | -0.019            | 0.041            |
|  | non-labour income          | -0.039       | 0.003            | 0.020             | 0.040            |
|  | net wage                   | 0.050        | 0.004            | -0.023            | 0.051            |
|  | transfer                   | -0.019       | 0.001            | 0.008             | 0.017            |

*Notes:* Column (1) reports probit estimates and conditional marginal effects computed from the estimation on the full sample and evaluated at the subgroup-specific mean values of the covariates. Column (2) reports similar marginal effects, but computed from the estimations on the restricted samples.

**Table 2.4.:** Conditional marginal effects by selected subgroups

|                                       |                   | <i>dy/dx</i> | <i>std. err.</i> |
|---------------------------------------|-------------------|--------------|------------------|
| elder (>=50)                          | gains to work     | 0.311        | 0.052            |
|                                       | non-labour income | -0.320       | 0.057            |
|                                       | net wage          | 0.392        | 0.065            |
|                                       | transfer          | -0.103       | 0.017            |
| women at<br>child-bearing age (25-49) | gains to work     | 0.146        | 0.013            |
|                                       | non-labour income | -0.151       | 0.014            |
|                                       | net wage          | 0.231        | 0.021            |
|                                       | transfer          | -0.108       | 0.010            |
| prime-age, single men                 | gains to work     | 0.069        | 0.008            |
|                                       | non-labour income | -0.071       | 0.009            |
|                                       | net wage          | 0.096        | 0.012            |
|                                       | transfer          | -0.038       | 0.005            |
| prime-age, single women               | gains to work     | 0.113        | 0.013            |
|                                       | non-labour income | -0.116       | 0.013            |
|                                       | net wage          | 0.168        | 0.019            |
|                                       | transfer          | -0.076       | 0.008            |
| prime-age, married men                | gains to work     | 0.028        | 0.003            |
|                                       | non-labour income | -0.029       | 0.004            |
|                                       | net wage          | 0.039        | 0.005            |
|                                       | transfer          | -0.016       | 0.002            |
| prime-age, married women              | gains to work     | 0.183        | 0.016            |
|                                       | non-labour income | -0.189       | 0.017            |
|                                       | net wage          | 0.290        | 0.025            |
|                                       | transfer          | -0.133       | 0.012            |

*Notes: The table reports conditional marginal effects computed from the estimation on the full sample and evaluated at the subgroup-specific mean values of the covariates.*

### 2.3.2.3. Economic significance of the estimated elasticities

The estimated model can be utilized for the simulation of the labour supply (participation) effect of a personal income tax and transfer reform. The main step is to calculate the probability of being active for a given hypothetical wage, tax and transfer system. First we obtain the pre- and

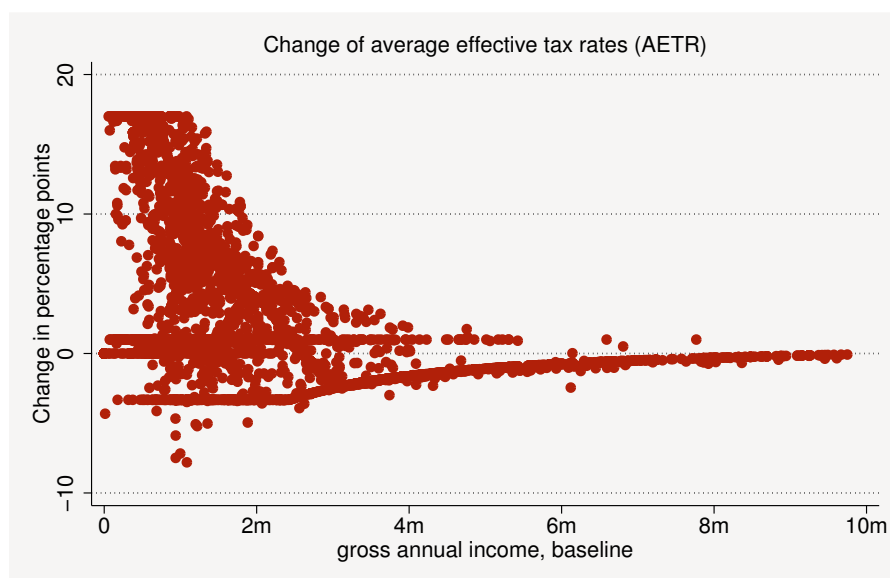
post-reform after-tax wage income of everyone in our sample, using predicted wages. Then we calculate the pre- and post-reform hypothetical “zero hours worked” transfer level for everyone, and construct  $\log W$  before and after the reform.

Equipped with these, we form

$$\Phi \left( \hat{\gamma} \log W_i + Z_i \hat{\theta}' - \hat{\psi} \log T_i \right)$$

before and after the reform. The change in its value is the change in the probability of individual  $i$  being active. Finally, we add up the probabilities in the sample (weighted) to get an estimate for the change in the aggregate activity rate. This gives us the shift of the labour supply curve: in equilibrium, labour demand might be downward sloping so the equilibrium wage may change, offsetting partly the change in labour supply.

**Figure 2.3.:** Change of average effective tax rates (AETR)



As an illustration, we fed the main changes of the Hungarian personal income tax system of 2012 into this framework. The particular measures are the following: the complete elimination of the employee tax credit (ETC) scheme, a reduction in the tax rate from 20,3% to 16% below the average monthly income of 202,000 HUF,<sup>18</sup> and a 1 percentage point increase in the social contribution rate. As illustrated by Figure 2.3, these changes have a very heterogeneous effect on the average tax rate of taxpayers: the abolishment of the ETC pushes up the average tax rate for low earners, for which they are partly compensated by the cut in the tax rate. Medium earners, who were not or at most

<sup>18</sup>The average exchange rate in 2012 was 290 HUF per euro.

partially eligible for the ETC gain by a reduction in their tax rate. High earners also gain a little due to the reduction in the tax rate on their first 202,000 HUF income per month. Finally, there is a common loss from increased social contributions.

As a result, aggregate activity decreases by 0.97 percentage points, from which the elimination of the ETC is responsible for 2.09 percentage points,<sup>19</sup> the increase in social contributions leads to another reduction of 0.34 percentage points, which are partly offset by an increase of 1.51 percentage points due to the rate cut.<sup>20</sup>

## 2.4. Conclusion

This paper provides an alternative modelling strategy of structural labour supply at the extensive margin to the two dominating approaches based on marginal calculus (Hausman (1981)) and on random utility models (van Soest (1995)). The “gains to work” model builds on the widely used discrete choice approach based on random utility models, in which utility-maximising individuals choose between discrete sets of hours worked based on their consumption-leisure preferences. Our model essentially requires the same assumptions. However, instead of explicitly parametrising the full underlying preference specification, we derive a labour supply function in which the individual compares the financial gains from accepting a job (the gains to work) with a certain threshold that can be interpreted as the “reservation gains to work”. The basic concept behind the resulting labour supply specification is similar to the labour supply models based on marginal calculus (Hausman (1981)), but it is adopted to reflect the nonlinearities of the choice set. In a sense, our model provides a bridge between the two dominating structural labour supply models in the literature, the early generation “continuous” Hausman labour supply model and the new discrete choice approach based on random utility functions.

We estimate the binary version of the model on Hungarian Household Budget Survey data. We find that a single equation can explain a large heterogeneity of individual responsiveness to taxes and transfers: there are marked differences in the conditional marginal effects among subgroups, driven partly by a composition effect (due to the nonlinearity of the probit function), and partly by a different share of lost transfers in the gains to work. The most responsive subgroups are exactly the ones who are mostly responsible for Hungary’s low participation rate (low-skilled, women at

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<sup>19</sup>There is a subtle issue here: under the Hungarian tax code, a large part of social transfers are also affected by personal income taxes and the ETC. Consequently, the elimination of the ETC also decreases the net value of many social transfers. Thanks to our integrated treatment of taxes and transfers, we can take this into account in our calculation. Without the corresponding cut in the net value of transfers, there would be an even more substantial reduction in participation.

<sup>20</sup>The sum of the effects of these measures may differ to the total effect due to interactions.



child-bearing age, elders), implying that a reform of the tax and transfer system can be a powerful tool to boost employment.

Our results directly lend themselves to reform simulations. We showed how our model can be utilized to calculate the labour supply shift of a complex personal income tax reform. In related work (Benczúr et al. (2012)), we build a model where this labour supply block is expanded by an intensive margin adjustment (based on a combination of Bakos et al. (2008) and Áron Kiss and Mosberger (2015)), and then it is embedded in a small general equilibrium macro model. With such a fully fledged model, we are able to evaluate at depth the 2011-12 Hungarian tax and transfer reforms as well (Benczúr et al. (2011) and Benczúr et al. (2012)).

# 3. Labour Force Participation and Tax-Benefit Systems: A Cross-Country Comparative Perspective

## 3.1. Introduction

Cross-country comparison of labour force participation rates is the most straightforward way to identify labour supply problems in a particular country or set of countries. International organisations, central banks and economic think-tanks often motivate their policy recommendations by comparing a country's participation rate — usually broken down by selected sub-populations — against an international standard or against other similar countries' statistics. The resulting recommendations frequently involve — but are not limited to — reform of the country's tax and welfare system to restore or improve work incentives.<sup>1</sup> That is, the financial incentive to work is implicitly the primary candidate for explaining cross-country differences in the participation rates of apparently homogeneous groups of individuals. Is this really the most important factor in explaining cross-country differences, or are other aspects such as differences in preferences or non-financial incentives even more important? To what extent can differences in financial incentives explain

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<sup>1</sup>For example, see OECD (2012), OECD (2014) and Fund (2013) for assessments of labour force participation in Hungary and the Czech Republic — the two countries this paper focuses on — and for resulting policy recommendations. By comparing labour force participation rates in OECD countries, OECD (2012) concludes that "... structural reforms are also needed (in Hungary) to better exploit existing resources and raise one of the lowest activity rates in the OECD". Similarly, Fund (2013) urges Hungary to "... raise the exceptionally low labor participation rate". OECD (2014) welcomes the fact that Czech "labour force participation has increased to the European average", although the relatively low labour force participation of women is identified as a bottleneck for the Czech Republic as well. In order to catch up with countries with higher participation rates, IMF staff recommend that Hungary implement "a more employment-friendly taxation for low income earners" and "raise women's participation in the labor market by reorienting public spending from cash benefits [...] towards the development of high quality early childhood education and day care centers." Similarly, the OECD recommends that Hungary "preserve work incentives when lowering the tax wedge" to "promote the participation of the elderly" and to "reform family policies to enhance women's labour market participation".

the observed differences between countries in labour force participation? Although the impact of labour taxation and the welfare benefit system on labour force participation has been widely investigated in the empirical literature, the extent to which cross-country differences in participation rates can be explained by the diversity of tax-benefit systems has never been addressed directly.

To fill the gap, we take the example of two countries, the Czech Republic and Hungary. We use, for both countries, a very detailed, complex microsimulation model and perfectly comparable micro estimates of labour supply at the extensive margin (i.e. the participation decision) to quantify the portion of the divergence in the two countries' participation rates explained by differences in their taxation and welfare benefit systems. More precisely, we first replicate for Czech individual-level data the labour supply estimation for Hungary presented in Benczúr et al. (2014). The two entirely comparable estimated equations for the Czech Republic and Hungary are then used to simulate how the aggregate participation rate would change in one country if it adopted the other country's tax and social welfare system.

To the best of our knowledge, this paper offers the first tentative explanation of cross-country differences in labour force participation rates by differences in taxation and welfare benefits. The closest strand of literature uses cross-country individual-level or disaggregated data either to investigate the labour supply effects of a specific policy change or to explain developments in participation rates from a cross-country comparative perspective. Without the sake of completeness, some recent cross-country empirical studies analyse the effect of hypothetical welfare reforms (e.g. Immervoll et al. (2007)); others study the effects of tax reforms (see e.g. Colombino et al. (2000), or Paulus and Peichl (2008), for a cross-country analysis of flat-tax reforms); cross-country microsimulation models are also used to estimate governments' redistributive preferences (Blundell et al. (2009); Bargain et al. (2011)). Some other papers use individual-level or disaggregated data to describe the evolution of labour supply in some selected countries, either for the whole population (Balleer et al. (2009), or Blundell et al. (2011)) or for a selected sub-population (e.g. Cipollone et al. (2013), for women). However, to the best of our knowledge, none of these papers or any other existing empirical ones have so far provided direct evidence to explain differences in participation rates between countries.

The Czech Republic and Hungary provide an interesting case for comparison. The two countries exhibit a lot of similarities: their economies are geographically close, are similar in size and level of economic development, and partly share a common history. In particular, both economies experienced full employment until the regime change at the end of the 1980s. Despite all their common factors, their participation rates differ markedly nowadays. In 2013, with only 57.5% of the 15–74 years old population working or actively seeking employment, Hungary recorded the fourth lowest participation rate among the EU member states (behind Croatia, Italy and Malta), while the Czech

participation rate was close to the EU15 average (64.3% in the Czech Republic and 64.8% in the EU15).<sup>2</sup>

These differences are typically explained by different labour market policies adopted during the first few years of the transition process and kept – at least partly – unchanged since then. Both countries implemented fast-track reforms in the early 1990s, with two main differences. First, Hungary opted for case-by-case privatisation as opposed to the voucher privatisation method adopted by the former Czechoslovakia. Second, the Hungarian government introduced a strict bankruptcy regulation in 1992.

After an initial collapse, the widespread changes to the economic system shifted both economies to a relatively fast growing path. Nevertheless, the differences in the privatisation methods and the draconian bankruptcy regulation implemented in Hungary provoked a much larger drop in employment in Hungary as compared to Czechoslovakia. Between 1990 and 1993, total employment declined by 9% in Czechoslovakia and by 22% in Hungary. Moreover, in Hungary, a number of policy measures – such as alleviated conditions for entering the old-age or disability pension systems – contributed to pushing people out of the labour force rather than into unemployment. As a result, the participation rate declined continuously until 1997, when it reached only 50.6% of the 15–74 years old population, about 10 percentage points lower than the EU15 average and more than 13 percentage points lower than in the Czech Republic.<sup>3</sup>

As we will see, in 2008, the Hungarian transfer system could still be viewed as “generous” relative to the Czech system. After presenting our methodology (in the next section), we describe the key differences between the two countries’ tax and social welfare systems (Section 3.3).

Our results (presented in Section 3.4) show, first, that the estimated labour supply elasticities for the Czech Republic are very close to the results for Hungary, suggesting that, at least in this dimension, individual preferences are similar in the two countries. This holds true even for sub-populations depending on the level of education, gender and marital status: in both countries, lower educated people, the elderly and married women (of childbearing age) are the most responsive to tax and transfer changes. Given the large dispersion of the labour supply elasticities found in the empirical literature for different countries and time periods, the similarities of the estimated elasticities for the two countries we analyse are in line with the evidence in Bargain et al. (2012) suggesting that a considerable part of the cross-country variation in the elasticities found in the literature is driven by methodological differences.

Second, the simulation results suggest that about one-half of the total difference in the participation rates of the 15–74 years old population can be explained by differences in the tax-benefit systems.

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<sup>2</sup>Eurostat data, September 2014.

<sup>3</sup>For a detailed analysis of the evolution of the Hungarian participation rate, see Kátay and Nobilis (2009).

The results are quasi-symmetric, meaning that if the Czech system was adopted, the Hungarian participation rate would increase by about the same number of percentage points that the Czech participation would decrease by if the Hungarian system were implemented. The highest responses are obtained for married women and women of childbearing age. This is related to the much more generous maternity benefit system in place in Hungary as compared to the Czech Republic.

## 3.2. Methodology

We first replicate the estimation in Benczúr et al. (2014) for Czech data and compare our results with the Hungarian estimates of labour supply at the extensive margin. By closely following the approach specified in Benczúr et al. (2014), we estimate a fully parametric structural labour supply model where both taxes and transfers are treated in a unified framework. In the second part, we use the two estimated equations for Hungary and the Czech Republic to simulate how each individual's probability of being active — and hence the aggregate labour supply — would change in one country if it adopted the other country's tax-benefit system. More precisely, we perform a microsimulation exercise: we replace the net income variables (wages, transfers and non-labour incomes) in the estimated equations by their values that would result from adopting the other country's tax and transfer system. As a consequence of the benefits gained or losses suffered due to changes in the effective tax rates and the amount of (potential) transfers, individuals are expected to adjust their labour supply according to the estimated probit equations. The weighted average of the individual changes in the participation probabilities corresponds to the aggregate labour supply shock induced by the hypothetical reform. Under certain assumptions detailed later, the result of this exercise reveals the extent to which the difference between the two countries' participation rates can be explained by the differences in the taxation and welfare benefit systems.

### 3.2.1. Modelling Labour Supply at the Extensive Margin

An individual's labour-supply decision is usually modelled as a trade-off between the utility gained from consumption or leisure. Starting from a standard utility function characterised by strictly positive marginal utilities with respect to consumption and leisure, the participation decision is determined by the difference between the market wage and the reservation wage. In the standard textbook approach based on marginal calculus, non-participation follows simply from the corner solution of the model (Hausman (1981)). The theory yields a binary dependent variable specification in which the probability of working or actively seeking work is modelled as a function of

the net market wage, net non-labour income and individual preferences. In this framework, taxes influence labour supply by affecting net market wages and net non-labour incomes.<sup>4</sup>

The early generation Hausman model proved to be too restrictive in many respects. First, relying on tangency conditions, the Hausman model is restricted to the case of (piecewise) linear and convex budget sets. This assumption is particularly restrictive if certain transfers get lost immediately on taking up a job — which is usually the case — and the wage earned during the first few hours worked does not compensate for the discrete loss in benefits. As a consequence, the standard reservation wage — defined as the lowest wage rate at which a worker would be willing to work nonzero hours — is undefined.<sup>5</sup> Second, quasi-concavity of the utility function is implicitly imposed a priori. As discussed by MaCurdy et al. (1990), the Hausman model requires the Slutsky condition to hold at all internal points of the budget constraint. This restriction is, however, typically rejected by empirical studies (see e.g. Blundell and MaCurdy (1999)) and may lead to biased estimates by implicitly limiting the range of elasticities that can be obtained. Third, predictions using the standard approach do not fit the data well, as the model does not account for the fact that very few observations exist with a small positive number of hours worked. Finally, the model makes it difficult to handle household members' joint labour supply decisions.

In response to these shortcomings, the discrete choice approach to modelling labour supply provides a convenient alternative to the continuous hours methodology. Originally proposed by van Soest (1995), this approach has become increasingly popular and quite standard in recent years. In this framework, utility-maximising individuals are supposed to choose between a few alternative discrete sets of hours of work, such as inactivity (zero hours worked), part-time or full-time. Reducing the maximisation problem to choosing among a discrete set of possibilities yielding different utilities considerably simplifies the problem and provides a simple yet rather general way of representing labour supply decisions in the presence of nonlinear and non-convex budget constraints.

Mathematically, the utility ( $U$ ) that individual (or household)  $i$  derives from choosing alternative  $j$  from  $J$  possible discrete choices is represented by a random utility model of the form:

$$U_{ij}(c_{ij}, 1 - l_j) = V(c_{ij}, 1 - l_j, Z_i, \theta) + \varepsilon_{ij} \quad (3.1)$$

where  $V(c_{ij}, 1 - l_j, Z_i, \theta) = V_{ij}$  is a positive deterministic term that represents the mean utility across observationally identical agents when alternative  $j$  is chosen and  $\varepsilon_{ij}$  is a random term that is not correlated with the deterministic part of the utility function.  $V_{ij}$  is conditional on the alternative-

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<sup>4</sup>For a comprehensive presentation of the different variants of the standard modelling approach and their identification strategies, see e.g. Blundell and MaCurdy (1999) and Blundell et al. (2007)

<sup>5</sup>To overcome this deficiency, fixed costs of work can be introduced, yet this additional source of non-convexity is also difficult to handle (see e.g. Bourguignon and Magnac (1990)). van Soest and Das (2001) argue that non-convexities in general imply rather restrictive and implausible forms for preferences.

specific consumption ( $c_{ij}$ ) and leisure ( $1 - l_j$ , with total time endowment normalised to 1), a set of individual characteristics ( $Z_i$ ) and preference parameters ( $\theta$ ).

In our empirical work, we model individual labour supply decisions. We assume that there are only two labour-market states, active and inactive. Indeed, in both the Czech Republic and Hungary, part-time work is relatively rare: before the outbreak of the current crisis, the share of part-time employees among all workers was less than 5% in both countries. By 2013, this share had increased to 6.6% in the Czech Republic and to 6.7% in Hungary. For comparison, the share of part-time employees in the EU15 reached 21% in 2008 and increased further to 23.6% in 2013.<sup>6</sup>

The participation decision boils down, therefore, to a comparison of the utility gained from working full-time ( $j = 1$ ) to the utility gained from staying out of the labour force and getting the full amount of transfers ( $j = 0$ ). Individual  $i$  therefore chooses to work full-time if  $U_{i1} \geq U_{i0}$ . Given the budget constraint stating that consumption must equal total income, the probability of participation is given by:

$$\begin{aligned} P_{i,j=1} &= P [U_{i1}(w_i + T_i - \Delta T_i, 1 - l_1) \geq U_{i0}(T_i, 1)] \\ &= P [V(w_i + T_i - \Delta T_i, 1 - l_1, Z_i, \theta) + \varepsilon_{i1} \geq V(T_i, 1, Z_i, \theta) + \varepsilon_{i0}] \end{aligned} \quad (3.2)$$

where  $P_{i,j=1}$  represents the probability that individual  $i$  is economically active;  $w_i$  being the net market wage of individual  $i$  if working full-time ( $l = l_1$ );  $\Delta T_i$  is the sum of all the hypothetical social transfers the individual gets (or would get) at zero hours worked and loses when working full-time; and  $T_i$  is the total net non-labour income, including social transfers ( $\Delta T_i$ ) and other non-labour income (pensions, dividend payments, income of other members of the household, etc.)

If the random components  $\varepsilon_{ij}$  are i.i.d. extreme value distributed with c.f.d.  $\exp\{-\exp(-\varepsilon_{ij})\}$  with fixed variance, McFadden (1974) derives a formal expression of the probability that alternative  $j$  is chosen and shows that the parameters can be estimated by maximum likelihood techniques.<sup>7</sup>

Following Benczúr et al. (2014), we take a somewhat different approach. Benczúr et al. (2014) transform eq. (3.2) by linearising the left-hand side of the expression:

$$U_{i1}(W_i + T_i, 1 - l_1) \approx U_{i1}(T_i, 1 - l_1) + W_i U'_{i1}(T_i, 1 - l_1) \quad (3.3)$$

<sup>6</sup>Source: Eurostat

<sup>7</sup>This model, however, exhibits the unpleasant property of Independence of Irrelevant Alternatives (IIA). To overcome this deficiency, empirical studies usually introduce some unobserved heterogeneity in preferences, that is, some of the parameters are assumed to be randomly distributed. The estimation technique for this model, referred to as “mixed logit” or “random parameter logit”, is described in detail in McFadden and Train (2000).

where  $W_i = w_i - \Delta T_i$  represents the “gains to work”, defined as the difference between the net wage and the amount of transfers lost when working. The comparison becomes:

$$W_i \geq \frac{U_{i0}(T_i, 1) - U_{i1}(T_i, 1 - l_1)}{U'_{i1}(T_i, 1 - l_1)} \quad (3.4)$$

The individual therefore chooses to participate if the gains from accepting a full-time job (the gains to work) exceeds a certain threshold, which can be interpreted as the “reservation gains to work”. This concept is similar to the standard textbook approach and can be interpreted as the discretised version of the Hausman method.

By log-linearising the right-hand side of eq. (3.4) and approximating the unobserved components by  $Z_i\alpha' + \varepsilon_{it}$  with  $\varepsilon_{it} \sim N(0, 1)$ , Benczúr et al. (2014) derive a formal expression for the probability of being active, which yields the following structural probit equation:

$$P_{i,j=1} = \Phi(\gamma \log W_i + Z_i\alpha' + \psi \log T_i) \quad (3.5)$$

In this framework, instead of directly estimating the parameter of the utility function, we approximate the “reservation gains to work” by using a set of observed individual characteristics ( $Z_i$ ) and the total net non-labour income ( $T_i$ ).

In practice, not all the variables of eq. (3.5) are observed for all individuals. As in Benczúr et al. (2014), we construct the missing variables as follows:

1. for employed individuals, we directly observe their gross wages, so we can compute their net wages ( $w_i$ ).<sup>8</sup> However, we do not observe the hypothetical amount of transfers the individual would get if he or she decided to stop working. Using individual characteristics and the details of the welfare system, we calculate the amount of lost transfers ( $\Delta T_i$ ) the individual would be entitled to being non-employed. Consequently, the gains to work are calculated as  $W_i = w_i - \Delta T_i$ , while the amount of lost transfers ( $\Delta T_i$ ) is added to the observed non-labour income to obtain  $T_i$ .
2. for the non-employed, the total amount of transfers is observed, whether income-tested or not.<sup>9</sup> We do not observe, however, the market wage. Similarly to Benczúr et al. (2014), the gains to work are estimated using the Heckman selection model (see Section 3.2.2).

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<sup>8</sup>Gross wages include employer contributions as well. Accordingly, we take into account both employer and employee-side contributions and taxes when computing net wages.

<sup>9</sup>We can also determine  $\Delta T_i$  the same way we computed it for employed individuals by applying the welfare rules. We used the observed values for the estimation part; however, we used our own calculation for the simulation part. See Section 3.2.3 for details.



### 3.2.2. Estimation Method

Equation (3.5) is estimated using a three-step procedure. The first two steps represent in fact a Heckman sample selection model to predict the gains to work. The only difference compared to the standard Heckman procedure is that instead of wages, we directly estimate the gains to work. The latter are supposed to be influenced by a vector  $X_i$ , which includes variables affected by both labour supply and labour demand factors. The gains-to-work equation is thus specified as:

$$\log W_i = X_i\beta' + \mu_i \quad (3.6)$$

The three-step procedure is thus as follows:

1. first, we run a reduced-form probit equation explaining the probability of being employed ( $e = 1$ ):

$$P_{i,e=1} = \Phi(X_i\beta'_{RF} + Z_i\alpha'_{RF} + \psi_{RF} \log T_i) \quad (3.7)$$

2. second, we use the inverse Mills ratio  $\lambda(x) = \phi(x)/\Phi(x)$  as a correction in the gains-to-work equation:

$$\log W_i = X_i\beta' + \lambda(X_i\beta'_{RF} + Z_i\alpha'_{RF} + \psi_{RF} \log T_i) + \mu_i \quad (3.8)$$

3. third, we use the predicted gains to work  $\log \hat{W}_i = X_i\hat{\beta}'$  to estimate the structural probit equation<sup>10</sup>:

$$P_{i,j=1} = \Phi(\gamma \log \hat{W}_i + Z_i\alpha' + \psi \log T_i) \quad (3.9)$$

To put it differently, the estimation procedure corresponds to a IV probit where the first step is a Heckman selection model. The endogeneity and measurement error biases associated with  $\log W_i$  are therefore corrected.<sup>11</sup> The key element of the identification is the presence of “labour demand shifters” in  $X_i$  which are missing from the structural equation and act like excluded instruments,

<sup>10</sup>Note that we used the unconditional prediction, that is, the effect of the inverse Mills ratio is not included.

<sup>11</sup>The wage error term may indeed be correlated with the participation error term. Note also that  $\log W$  is subject to measurement error for various reasons. For example, observed wages are biased downward for part-time workers or for individuals who work less than ~20 days per month. Lost transfers ( $\Delta T$ ) – which are part of  $\log W$  – may also be mismeasured due to possible lower take-up rates of some benefits and due to the fact that some transfers may be provided with discretion by social offices. It is possible, however, to compare the simulated  $\Delta T$  with the observed transfers for individuals who worked for some part of the year and were non-employed for the rest of the year, as both their wages and their transfers are observed. This comparison revealed that for both countries and for most individuals, the simulated transfers were close to the observed ones.

that is, variables which determine the market wage but only affect labour supply indirectly through wages.

The labour supply shifters ( $Z_i$ ) include education dummies, household head, mother with infant (< 3 years old), attending full-time education, household size (number of persons), pensioner, family status (husband, wife, child, single, divorced, etc.) and age-group (15–24, 25–49, 50+) dummies and year dummies.

Following Benczúr et al. (2014), the labour demand shifters ( $X_i$   $Z_i$ ) contain age, age squared and education dummies and their interactions, and county dummies. The county dummies are proxies for local labour market conditions, while age and age square and their interactions with the education dummies are proxies for experience that influence market wages. Once we control for individuals' lifecycle position (student, family status, age group, retired, etc., all variables included in  $Z_i$ ), we assume — in line with Benczúr et al. (2014) — that an extra year has a negligible impact on labour supply directly, but it strongly affects the wage and hence influences labour supply indirectly.

Once estimated, the structural probit estimation gives us the coefficients of  $\log W$  and  $\log T$  ( $\gamma$  and  $\psi$  respectively) as well as the conditional marginal effects of these variables — evaluated at a certain vector, e.g. at the sample means — on the probability of being active. To obtain the impact of the net wage ( $w$ ) itself, Benczúr et al. (2014) show that the marginal effect of has to be multiplied by  $w/(w - \Delta T)$ . Similarly, the effect of conditional transfers on activity is given by  $(\partial\Phi/\partial\log W)(1 - w/(w - \Delta T)) + (\partial\Phi/\partial\log T)(\Delta T/T)$ . We report these marginal effects in Section 3.4.

### 3.2.3. Simulations

The simulations are carried out symmetrically using both the Hungarian and the Czech results. As a first step, we predict the participation probabilities using eq. (3.5) estimated for the Hungarian and Czech data as follows:

$$\begin{aligned} \hat{P}_{i,j=1}^{CZ,CZ} &= \Phi \left( \hat{\gamma}^{CZ} \log W_i^{CZ} + Z_i \hat{\alpha}^{CZ} + \hat{\psi}^{CZ} \log T_i^{CZ} \right) \\ \hat{P}_{i,j=1}^{HU,HU} &= \Phi \left( \hat{\gamma}^{HU} \log W_i^{HU} + Z_i \hat{\alpha}^{HU} + \hat{\psi}^{CZ} \log T_i^{CZ} \right) \end{aligned} \quad (3.10)$$

$\hat{P}_{i,j=1}^{C_1,C_2}$  denotes the predicted probability of individual  $i$  in country  $C_1 = \{CZ, HU\}$  being active and using the tax and transfer system of country  $C_2 = \{CZ, HU\}$ . Parameters with superscripts  $CZ$  and  $HU$  are estimated parameters for the Czech and Hungarian data respectively.  $W_i^{C_1}$  and  $T_i^{C_1}$  are the gains-to-work and the non-labour income (including conditional transfers) variables

calculated using country  $C_1$ 's tax and transfer system. For example, in the first equation of 3.10, we took the observed gross wages for employed Czech individuals and the estimated gross market wages for non-employed individuals and we applied the Czech tax-benefit system to compute  $W_i^{CZ}$  and  $T_i^{CZ}$  for each individual. In eq. 3.10, we set the constant term so that the weighted average of the predicted probabilities exactly matches the aggregate participation rate for the whole population.

In addition to labour income taxes, we took into account indirect taxes (that is, the average effective VAT rate, denoted as  $\nu$ ) in our simulation exercise. The gains to work therefore become  $W = (w - \Delta T)/(1 - \nu)$ . Similarly, non-labour incomes are also divided by  $(1 - \nu)$ .<sup>12</sup>

As a second step, we convert all income variables in the Czech database to Hungarian forint (HUF) using the 2008 average exchange rate (CZK 1 = HUF 8.04).<sup>13</sup> Then we apply the Hungarian tax-benefit system to the Czech data to compute net income variables and thus to obtain  $W_i^{HU}$  and  $T_i^{HU}$ . Before calculating the implied predicted probabilities of being active, we convert the variables back to CZK. We perform symmetrically the same exercise by applying the Czech tax-benefit system to the Hungarian data. The predicted probabilities are given by:

$$\begin{aligned} \hat{P}_{i,j=1}^{CZ,HU} &= \Phi \left( \hat{\gamma}^{CZ} \log W_i^{HU} + Z_i \hat{\alpha}^{CZ} + \hat{\psi}^{CZ} \log T_i^{HU} \right) \\ \hat{P}_{i,j=1}^{HU,CZ} &= \Phi \left( \hat{\gamma}^{HU} \log W_i^{CZ} + Z_i \hat{\alpha}^{HU} + \hat{\psi}^{HU} \log T_i^{CZ} \right) \end{aligned} \quad (3.11)$$

The differences between these predicted probabilities and those obtained in the previous step — that is,  $\hat{P}_{i,j=1}^{CZ,HU} - \hat{P}_{i,j=1}^{CZ,CZ}$  and  $\hat{P}_{i,j=1}^{HU,CZ} - \hat{P}_{i,j=1}^{HU,HU}$  — correspond to the individual labour supply shocks induced in one country by adopting the other country's tax-benefit system. The weighted averages of these expressions give the aggregate labour supply shocks triggered by the hypothetical reforms.

Whether the initial labour supply shock fully translates into participation in the long run depends on the general equilibrium interactions of the entire economy. For example, in a textbook neoclassical general equilibrium long-run model of a small open economy, the overall effect primarily depends on the assumption made on the long-run elasticity of capital supply. In the perfectly elastic case — which is a standard assumption in the literature — the expected rate of return on capital is pinned down by international benchmarks, which implies that the capital-to-labour ratio and the ratio of factor prices stay constant in the long run. It follows that, for example, after a positive labour supply shock induced by a tax cut, capital accumulation will follow until the new equilibrium is

<sup>12</sup>In the estimation part, the effect of VAT is captured by year dummies.

<sup>13</sup>Alternatively, measures that also reflect differences in economic development, such as the ratio of the two countries' nominal GDP per capita, could be used. We also performed the simulations using this alternative measure, but do not present them in this paper. The results — available from the authors upon request — changed only marginally.

reached with an unchanged capital-labour ratio. After the adjustment period, gross wages will also reach their pre-reform level, so factor prices remain unchanged in the long run. In this framework, the initial labour supply shock fully translates in the long run into participation. Consequently, under these assumptions, the results of our simulation exercises can be interpreted as the effect of differences in the tax-benefit systems in the two countries on the differences in the two countries' participation rates.<sup>14</sup>

## 3.3. Key Elements of the Czech and Hungarian Tax and Transfer Systems

The Czech and Hungarian tax-benefit systems are both characterised by individual personal income taxation and a broad range of welfare benefits. In what follows, we briefly describe and compare the two systems effective in 2008, the reference year we used for our simulation. In this section, we only consider the main characteristics of the systems that apply to regular wage earners; the self-employed and other specific cases are not considered. In our empirical work, more details are taken into consideration. For a more comprehensive description of the two systems, see, for example, the country chapters of the OECD's Benefits and Wages publication.<sup>15</sup> For a detailed presentation of the calculation methods for net incomes and (hypothetical) transfers, see Benczúr et al. (2014) for the Hungarian case and Galuscak and Pavel (2012) for the Czech case. Additional details on the transfer systems of the two countries are presented in the Appendix.

In both countries, labour incomes are subject to personal income tax and employer and employee contributions. In Hungary, the tax schedule comprises three brackets, with tax rates of 18%, 36% and 40%.<sup>16</sup> Employees with annual earnings under a given threshold are eligible for an earning income tax credit (EITC) of 18% of their wage income, which is phased out at a rate of 9%. A child tax credit (fixed amount) is available for families with three or more children. Social security contributions related to sickness, unemployment and pensions are applied.

In the Czech Republic, family taxation for married couples with children was introduced in 2005. The joint taxation, however, became irrelevant when a flat tax rate of 15% was introduced in 2008. It is applied to so-called super-gross income, which includes employer social security contributions.

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<sup>14</sup>For a detailed explanation of the long-run general equilibrium effects of a tax reform and the implications of capital supply elasticity for the new equilibrium, see a closely related paper: Benczúr et al. (2012).

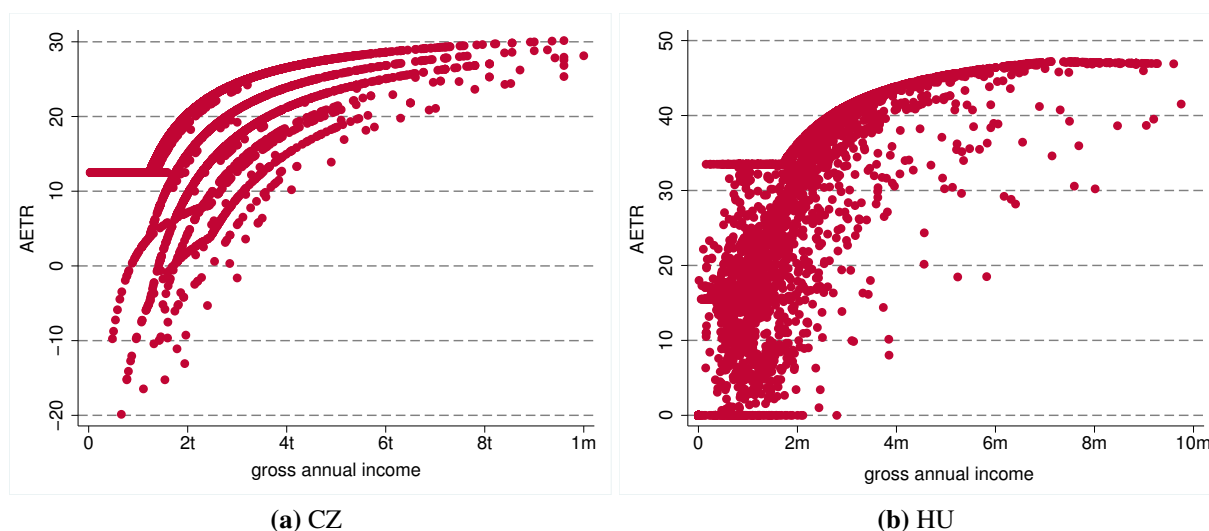
<sup>15</sup>[www.oecd.org/els/social/workincentives](http://www.oecd.org/els/social/workincentives)

<sup>16</sup>The third bracket is in fact the second bracket plus a 4% surtax -- a.k.a. the "solidarity levy" -- which was introduced in 2007.

Tax credits are available per person, per spouse under a given income and per dependent child. If the tax credit per child is negative, it is provided to the household as a tax bonus.

Comparing the two personal income tax (PIT) systems, Figure 3.1 reveals that, first, the Hungarian system is more progressive and, second, average taxes are higher, especially — but not only — for higher income earners. Moreover, the Czech average tax rate can be negative for families with children if the calculated income tax is lower than the amount of family benefit the taxpayer is eligible for. The higher average PIT rates in Hungary are also reflected in higher tax revenues: the Hungarian government in 2008 collected 7.9% of GDP as PIT tax revenue, compared to 3.7% in the Czech Republic (see Table 3.1).

**Figure 3.1.:** Average Earning Tax Rate (AETR) in CZ and HU (2008)\*



*Note: The average earning tax rate is the total amount of tax paid on labour earnings as a percentage of the tax base if all forms of income taxes, contributions, tax credits and bonuses are taken into account.*

Both countries provide a wide range of social benefits in order to — at least partly — compensate for temporary loss of labour income, to reduce and prevent poverty or to achieve other goals. Overall, if we look at the share of GDP spent on various transfers in 2008, the Hungarian benefit system seems more generous than the Czech one (Table 3.1).<sup>17</sup>

<sup>17</sup>The Hungarian benefit system is often cited as being relatively generous compared to the systems of neighbouring countries based on the share of GDP redistributed, on the eligibility criteria for some benefits and on the amount of welfare transfers an individual might be eligible for. At the same time, Hungary is not successful in ensuring access to transfers by the most vulnerable. In comparison with the Czech Republic, the share of people at risk of poverty or social exclusion is more than two times higher in Hungary than in the Czech Republic according to Eurostat data (33.5% vs 14.6% in 2013). The Hungarian redistribution is arguably far from the egalitarian principle.

### 3.3 Key Elements of the Czech and Hungarian Tax and Transfer Systems

**Table 3.1.: Key Differences in the Tax and Transfer Systems (2008)**

|                       | Description  |   | % of GDP |       |
|-----------------------|--|---|----------|-------|
|                       | CZ   | HU  | CZ       | HU    |
| Taxes on labour       |  |   |          |       |
| PIT                   | one tax rate (15% on supergross) + allowances (children), negative AETR possible   | three tax rates (18%, 36%, 40%) with ETC + allowances, no negative AETR   | 3.7%     | 7.9%  |
| employee contr.       | 12.5%  | 17%   | 3.5%     | 3.2%  |
| employer contr.       | 35%  | 32% + 1950 HUF/month  | 9.9%     | 9.9%  |
| Other taxes           |  |   |          |       |
| eff. tax on cons.     | 15.4%  | 18.2%   | 10.6%    | 15.6% |
| eff. tax on sales     | –  | 1.7%  | –        | 1.7%  |
| eff. tax on capital   | 9.8%   | 7.3%  | 4.2%     | 2.6%  |
| Benefits              |  |   |          |       |
| maternity benefit     | 6 months: 69% of previous gross income (capped)  | first 2 years: 70% of previous income (capped), third year: fixed amount; longer entitlement period for disabled child, twins or three children                       | 0.2%     | 0.7%  |
| UB                    | first 3 months: 50% of previous net income; next 3 months: 45% of previous net income. Capped; longer entitlement period for the elderly | max 3 months: 60% of previous income (capped); then max 6 months: fixed amount; then max 3 more months: lower fixed amount; longer entitlement period for the elderly | 0.2%     | 0.4%  |
| other social benefits | fixed amount   | fixed amount  | 1.2%     | 1.2%  |
| old-age pension       | official: 61 years and 10 months for men, 60 years for women (less with children); effective: 62.6 for men, 58.8 for women               | official: 62 years; effective: 59.5 for men, 58.5 for women   | 5.80%    | 9.7%  |
| disability pension    |  |   |          | 2.4%  |

Source: National Accounts (HU and CZ), ESCB Public Finance Report (2014), Czech Ministry of Labour. Notes: The effective tax on consumption is the sum of indirect taxes divided by final consumption expenditure; the effective tax on capital is corporate tax revenues divided by the share of capital revenues based on the National Accounts. Other social benefits include: (CZ) total state social support (including parental allowance and birth grant) and material need benefits and (HU) regular social assistance benefit, social supplement, other benefits (e.g. retraining), care allowances and other irregular social benefits.

As a general rule, in Hungary, maternity benefits and unemployment benefits either are fixed amounts or depend on previous income and are subject to income taxation. Other social benefits and pensions are, on the other hand, exempt from taxation. In the Czech Republic, all benefits are non-taxable (except very high pensions). In the case of maternity benefits, the amount is calculated based on previous *gross income*, whereas unemployment benefits are calculated based on previous *net income*. These differences make direct comparison of the two systems rather challenging. Some details of the two welfare benefit systems are listed in Table 3.1. For a more detailed description, see the Appendix or the references presented in the first paragraph of this section.

### 3.4. Results

The estimations are carried out using SILC (“*Survey of Income and Living Conditions*”) data for the Czech Republic for the years 2005–2010. The SILC is a yearly cross-section survey of households which provides detailed information on demography and incomes. For each year, the survey contains broadly 10,000 households and 23,000 individuals.

All income variables (labour and non-labour incomes, transfers) are expressed in total annual terms. To compute net incomes and potential transfers that individuals are entitled to or would be entitled to in the absence of any income from work, we used a modified version of the tax-benefit model developed by Galuscak and Pavel (2012). We used the same codes to compute net labour incomes, but we extended the original version of the model in order to take into account more complex features of the Czech tax and transfer system (e.g. taxes on capital revenues, maternity benefit, etc.). As for the transfers, we simulated the total yearly amount of benefits the individual gets if he is not working or could get if he decided to stop working.<sup>18</sup> At the end, we obtained the same model as in Benczúr et al. (2014) for Hungarian data.

The estimation results reported for Hungary are from Benczúr et al. (2014).<sup>19</sup> The Hungarian Household Budget Survey (HBS) is more detailed but otherwise perfectly comparable to the CZ-SILC database. First, several income variables in the Czech database are broken down into several categories in its Hungarian counterpart. Second, some variables — such as income from abroad, income of minors, life annuity payments, scholar grants and severance indemnities — are reported in the Hungarian database but no information is available from the Czech database. For the simulation exercise, we carefully linked all the variables to their counterparts in the other database. If

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<sup>18</sup>The original tax-benefit model in Galuscak and Pavel (2012) instead simulates the amount of monthly benefits an individual can get at the moment of the survey.

<sup>19</sup>The estimations for Hungary were carried out using the HBS database for the years 1998–2008.

one-to-one correspondence could not be achieved, the closest possible several-to-one match was established.

### 3.4.1. Estimation Results

Table 3.2 reports the baseline estimates, following the methodology presented in Section 3.2.2. As expected, higher gains to work increase the probability of participation in the labour market, while non-labour income has the opposite effect.

Overall, the results are remarkably close to the Hungarian ones. The point estimates and the conditional marginal effects of gains to work — evaluated at the population averages — are basically the same in the two countries. This is also true for the conditional marginal effect of net wages, suggesting that, on average, a similar change in the average tax rate has the same effect in both countries.

As for non-labour income, the estimated coefficient and the marginal effect are lower in absolute value in the Czech Republic than in Hungary, but the results for the conditional marginal effect of transfers are very close to each other. This seemingly confusing result is a reflection of the more detailed non-labour income (independent of working status) data used for the Hungarian estimation. As explained in Section 3.2.2, the marginal effect of conditional transfers ( $\Delta T$ ) is obtained by using the following formula:  $(\partial\Phi/\partial\log W)(1-w/(w-\Delta T))+(\partial\Phi/\partial\log T)(\Delta T/T)$ . Given that the Hungarian non-labour income contains more information ( $T$  includes certain types of incomes missing from the Czech database), the share of the potentially lost transfers in total non-labour income ( $\Delta T/T$ ) is higher in the Czech database. As a consequence, the (lower, in absolute value) Czech result for the marginal effect of non-labour income is multiplied by a relatively higher positive value, which leads to a similar estimate for the conditional marginal effect of transfers.

We are ultimately interested in the effects of tax and transfer changes, that is, the conditional marginal effects of the net wage and transfers. On average, a 10% increase in the net wage leads to a 3.7 percentage points increase in the probability of being active in the Czech Republic and a 3.9 percentage points increase in Hungary. A 10% increase in transfers decreases the average individual's probability of being active by 1.2 percentage points in the Czech Republic and by about 1.4 percentage points in Hungary. These strikingly similar results suggest that individual preferences are quasi-identical in the two countries.



**Table 3.2.: Main Results**

|                   | (A) Estimation results           |           |        |           |
|-------------------|----------------------------------|-----------|--------|-----------|
|                   | CZ                               |           | HU     |           |
|                   | (1)                              |           | (2)    |           |
|                   | coef.                            | std. err. | coef.  | std. err. |
| gains to work     | 0.818                            | 0.254     | 0.82   | 0.099     |
| non-labour income | -0.301                           | 0.038     | -0.844 | 0.11      |
|                   | (B) Conditional marginal effects |           |        |           |
|                   | dy/dx                            | std. err. | dy/dx  | std. err. |
| gains to work     | 0.268                            | 0.071     | 0.29   | 0.028     |
| non-labour income | -0.099                           | 0.013     | -0.298 | 0.03      |
| net wage          | 0.373                            | 0.099     | 0.395  | 0.038     |
| transfer          | -0.121                           | 0.028     | -0.136 | 0.013     |

*Source: CZ-SILC database, 2005–2010. The Hungarian results are from Benczúr et al. (2014). Notes: Three-step estimates, as described in the paper, using a sample of individuals aged 15–74. Standard errors are bootstrapped with 200 replications. The structural probit equation includes: log of gains to work, log of non-labour income, mother with infant (less than 3 years old), full-time student and education (less than elementary school, elementary school, vocational, secondary education, tertiary education) dummies, age-group dummies (15–24, 25–49, 50+), pensioner, gender and head of household dummies, household size and family status dummies (single, married living together, married living separately, widow(er), divorced), household membership status dummies (husband, wife, companion, single parent, child, ascendant, other relation, non-relation, single), year dummies. The controls included in the reduced-form probit and the wage equation which are missing from the structural probit are: county dummies, interaction of age and age square with education dummies.*

Following Benczúr et al. (2014), we report in Table 3.3 the conditional marginal effects by selected subgroups depending on level of education, gender and marital status. Again, the results are very close to the Hungarian ones. Panel A shows that the responsiveness of participation to changes in both net wages and transfers is lower for persons with higher (tertiary) education. Similarly to the results reported in Benczúr et al. (2014), the conditional marginal effects calculated on the prime-

age population averages (25–54 years old, reported in Panel B) are lower (in absolute value) than the results for the whole population, indicating that prime-age individuals are less responsive to tax and transfer changes than the rest of the population. Within the prime-age population, however, the education level strongly influences individuals' responsiveness: the lower educated ("elementary school or less"), probably highly transfer-dependent, group is highly responsive, while the estimated elasticities for secondary and tertiary educated people are much smaller.

The remaining rows in panel B suggest that women's participation is more responsive than that of men, particularly for married women and women of child-bearing age. Finally, the results in panel C show that the reaction of elderly persons to changes in the net wage and transfers is even higher.

The estimation results and the implied wage elasticities are qualitatively in line with existing results in the empirical literature. However, direct comparison of the elasticities is not straightforward. There is a large consensus in the literature that women (and especially married women) respond more to changes in their net market wage than men. Lower educated and elderly individuals are also usually found to be more responsive than the average. The variation in the magnitude of the estimated labour supply elasticities found in the literature is nonetheless considerable and these reported elasticities are rarely directly comparable for several reasons. First, with some exceptions, early labour supply models using the continuous Hausman approach usually do not distinguish between the decision to participate (the extensive margin) and the decision regarding hours worked (the intensive margin) and only report uncompensated and/or compensated (Hicksian) hours elasticities, not the participation elasticity.<sup>20</sup> Second, although the discrete choice approach always includes, by construction, the participation elasticity, estimates from various studies are not easily comparable due to differences in specifications and definitions. A recent comprehensive comparison of the international evidence on labour supply elasticities by Bargain and Peichl (2013) shows that the large variance in the reported elasticities is — at least partly — explained by differences in specifications (most importantly, whether wages are predicted for all individuals or only for non-workers) and the time period considered. A meta-analysis by Chetty et al. (2013) provides a "consensus" elasticity: according to the authors, existing micro and macro evidence points towards an aggregate steady-state elasticity of labour supply at the extensive margin of 0.25.

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<sup>20</sup>For a comparative overview of the hours elasticities stemming from the Hausman approach, see e.g. Pencavel (1987) for married men, Killingsworth and Heckman (1987) for married women, or the more recent survey by Evers et al. (2008). Other literature surveys include Blundell and MaCurdy (1999), Blundell et al. (2007), Keane (2011) and Meghir and Phillips (2011), which focus mostly on analyses based on the standard continuous approach, but also include some more recent evidence from discrete choice methodology.

**Table 3.3.:** Conditional Marginal Effects by Selected Subgroups

|                                    |          | CZ     |           | HU     |           |
|------------------------------------|----------|--------|-----------|--------|-----------|
|                                    |          | dy/dx  | std. err. | dy/dx  | std. err. |
| (A) Full sample                    |          |        |           |        |           |
| all individuals                    | net wage | 0.373  | 0.099     | 0.395  | 0.038     |
|                                    | transfer | -0.121 | 0.028     | -0.136 | 0.013     |
| elementary school or less          | net wage | 0.231  | 0.097     | 0.294  | 0.089     |
|                                    | transfer | -0.077 | 0.03      | -0.093 | 0.028     |
| secondary education                | net wage | 0.307  | 0.079     | 0.31   | 0.031     |
|                                    | transfer | -0.102 | 0.023     | -0.118 | 0.012     |
| tertiary education                 | net wage | 0.157  | 0.032     | 0.139  | 0.015     |
|                                    | transfer | -0.043 | 0.007     | -0.045 | 0.005     |
| (B) Prime-age subsample            |          |        |           |        |           |
| prime-age subsample                | net wage | 0.099  | 0.023     | 0.127  | 0.014     |
|                                    | transfer | -0.028 | 0.005     | -0.054 | 0.006     |
| elementary school or less          | net wage | 0.419  | 0.124     | 0.409  | 0.04      |
|                                    | transfer | -0.149 | 0.039     | -0.194 | 0.019     |
| secondary education                | net wage | 0.091  | 0.023     | 0.122  | 0.012     |
|                                    | transfer | -0.026 | 0.005     | -0.054 | 0.005     |
| tertiary education                 | net wage | 0.073  | 0.015     | 0.05   | 0.004     |
|                                    | transfer | -0.019 | 0.003     | -0.019 | 0.001     |
| single men                         | net wage | 0.078  | 0.019     | 0.096  | 0.012     |
|                                    | transfer | -0.025 | 0.005     | -0.038 | 0.005     |
| single women                       | net wage | 0.212  | 0.058     | 0.168  | 0.019     |
|                                    | transfer | -0.064 | 0.015     | -0.076 | 0.008     |
| married men                        | net wage | 0.016  | 0.002     | 0.039  | 0.005     |
|                                    | transfer | -0.004 | 0.000     | -0.016 | 0.002     |
| married women                      | net wage | 0.236  | 0.067     | 0.29   | 0.025     |
|                                    | transfer | -0.06  | 0.015     | -0.133 | 0.012     |
| women of child-bearing age (25-49) | net wage | 0.266  | 0.078     | 0.231  | 0.021     |
|                                    | transfer | -0.069 | 0.018     | -0.108 | 0.01      |
| (C) The elderly                    |          |        |           |        |           |
| elderly (50+)                      | net wage | 0.524  | 0.155     | 0.392  | 0.065     |
|                                    | transfer | -0.227 | 0.06      | -0.103 | 0.017     |

Notes: Conditional marginal effects computed using the estimation on the full sample and evaluated at the subgroup-specific mean values of the covariates; Hungarian results are from Benczúr et al. (2014).

Overall, the marginal effects obtained for the Czech Republic and Hungary lie in the (rather large) range obtained in previous studies, especially studies using the discrete choice approach.<sup>21</sup> Given the large dispersion of the labour supply elasticities estimated for different countries and time periods and using various methodologies, the similarities of the results we obtain for two different countries is in itself interesting and we tentatively suggest that methodological differences might be responsible for a considerable part of the cross-country variation in the elasticities found in the literature. The findings of Bargain et al. (2012), who estimate a comparable labour supply equation for a large set of countries and show that cross-country differences in genuine work preferences are rather small, corroborate this view.

Finally, to produce an accurate comparison with previous findings, and using the “consensus” elasticity reported by Chetty et al. (2013), Benczúr et al. (2014) simulated the effect of a one per cent increase in the net wage on the aggregate participation probability using the Hungarian data and estimated equation.<sup>22</sup> The resulting aggregate elasticity of 0.28 is quite in line with the consensus.

### 3.4.2. Simulation Results

The simulations were carried out as explained in Section 3.2.3. The results are shown in Table 3.4. The figures reported indicate the participation rates for 2008 (the reference year chosen for our simulations)<sup>23</sup> for both countries on the overall sample of individuals aged 15 to 74 (rows (a) to (g)), for the prime-age subsample (rows (h) to (m)) and for elderly people (row (n)).

The participation rates are considerably higher in the Czech Republic than in Hungary according to official Eurostat statistics (see columns (1) and (6)). Columns (2) and (7) show that the simple aggregate participation rates obtained from our datasets are very close to the official Eurostat data. The difference in participation rates is 8.5 percentage points according to the Eurostat data and 6.6 percentage points in our datasets.

The simulation results using each country’s own tax and benefit systems are presented in columns (3) and (8). The reported numbers are weighted averages of the participation probabilities predicted by eq. 3.10. As mentioned in Section 3.2.3, the constant terms in eq. 3.10 are set so as to match the aggregate statistics. The simulated full-sample aggregate participations (cells (a3) and (a8))

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<sup>21</sup>As noted in Bargain and Peichl (2013), the traditional continuous approach seems to overstate the elasticities compared to discrete choice models. For example, the own-wage elasticities for married women are relatively large using the continuous technique (in several cases, larger than 1), whereas the recent discrete choice approach points towards smaller elasticities (ranging between 0.1 and 0.5 with some exceptions).

<sup>22</sup>The conditional marginal effects presented in Tables 2 and 3 indicate the effect of a one per cent increase in the net wage on the “average individual’s” participation probability, as opposed to the change in the aggregate participation probability following a one per cent change in the net wage as reported by Chetty et al. (2013).

<sup>23</sup>2008 is the last year in our database not (or little) influenced by the crisis.

are therefore equal to the observed participation rates (cells (a2) and (a7)). The differences in the participation rates for the selected subgroups (columns (3) and (8)) come from differences in preferences (captured by the labour supply controls in eq. 3.10), differences in productivity and thus in gross wages, differences in average tax rates, which influence net wages, and/or differences in potential transfers. For both countries, the simulated participation rates for the selected subgroups are close to the observed ones. This suggests that overall, the estimated equations capture individuals' heterogeneity quite well.

The predicted aggregate participation rates using eq. 3.11 are reported in columns (4) and (9). In columns (5) and (10) we show how Czech and Hungarian labour force participation would change (in percentage points) if one country adopted the other country's tax-benefit system. The results of our simulation exercise show that, overall, the Czech participation rate would decrease by 2.9 percentage points if the country adopted the Hungarian system, while the Hungarian participation rate would increase by 3.0 percentage points if the Czech system were adopted. These results are highly symmetric and indicate that the differences in the two tax-benefit systems explain about half of the total difference in the two countries' participation rates. The difference is even greater for prime-age individuals, reaching -3.3 percentage points for the Czech data and +3.9 percentage points for the Hungarian data (see row (h)).

The simulation results remain broadly symmetric for most of the selected subgroups. Note that nothing guarantees that the two simulations should be perfectly symmetric for all subgroups. The main reason for the divergence is that these subgroups are not perfectly homogeneous and so the composition of the workforce within the subgroup influences the results. The largest difference is obtained for the elderly: the change in the aggregate participation probability for Hungary is twice the change in the probability for the Czech Republic when the other country's system is adopted. This divergence is at least partly explained by the higher share of pensioners, most importantly disability pensioners (see Table 3.5). As explained in Section 3.2.2, the same percentage change in the net wage results in a larger percentage change in the gains to work if the amount of (potentially) lost transfers is relatively high. As the gains-to-work change is larger on average, the participation response is also larger in the case of Hungary. The share of lower educated elderly people is also much higher in Hungary than in the Czech Republic (28% vs 17%; see Table 3.5). As lower educated individuals are more responsive to tax and transfer changes, the change in the aggregate participation rate is larger for Hungarian than for Czech elderly people. The differences in the share of those receiving disability pensions and the share of lower educated individuals are valid for almost all subgroups and thus influence individuals' responsiveness to tax and transfer changes.

Although the responsiveness to changes in the tax-benefit system decreases with education level, rows (b) to (d) of Table 3.4 do not reveal significant differences in labour supply responses across

various educational groups. Moreover, the 2.4–2.8 percentage points difference in the participation rates of the lowest educated individuals explained by the differences in the tax-benefit systems seemingly contradicts the aggregate statistics, which do not support such a difference in favour of the Czech Republic: according to our data, the participation rate is about 2 percentage points higher in Hungary than in the Czech Republic, whereas there is no difference according to the official Eurostat data.

Again, the difference in the composition of the subgroups provides a remedy to the apparent contradictions. In the Czech Republic, the share of full-time students is higher in the lower educated subgroup (column (1), row (b) of Table 3.5). Full-time students are economically inactive and are weakly responsive to changes in the tax or welfare benefit system. Once we filter out full-time students, the Czech participation rate of lower educated people is 9 percentage points higher than the Hungarian one. With no full-time students included, rows (e) to (g) of Table 3.4 show that the difference in participation explained by our model is higher for lower educated people (3.6pp–3.7pp) than for individuals with secondary education (3.1pp–3.3pp) and for those with tertiary education (2.8pp). These differences reflect the differences in responsiveness to tax and transfer changes.

The highest response is obtained for women, most importantly for married women and/or women of childbearing age (rows (j), (l) and (m)). The large response of these subgroups is most probably related to the very generous maternity benefit system in place in Hungary as compared to the Czech Republic. Although the net amount of the benefit is usually higher in the Czech Republic during the first 6 months after childbirth, the entitlement period is much longer in Hungary: conditional on past employment, Hungarian mothers can receive up to 70% of their previous income until the second birthday of the youngest child and parents are eligible for a fixed amount until the third birthday of the youngest child. Even longer entitlement periods apply for the third child, for twins and for disabled children.

Finally, note that the other, “unexplained” half of the difference between the two countries’ participation rates is also likely to be affected by some elements of the transfer system that we do not explicitly model, most importantly the differences in the old-age and disability pension schemes. Although both types of pensions are included in the model, neither individuals’ retirement decision, nor their access to the disability pension scheme is taken into account in our microsimulation exercise. That is, we assume that individuals’ retirement decision, their access to the disability pension scheme and the amount of these pensions remain unchanged when the other country’s welfare benefit system is adopted. Although several restrictive measures have been adopted to reduce abuse of the disability pension scheme and to increase the effective retirement age, the shares of early-retired and disabled individuals are still relatively high in Hungary by international comparison.

Table 3.4.: Simulation Results (2008)

|  | Czech Republic (CZ) |                   |                     |                 |               | Hungary (HU)      |                     |            |               |     |
|--|---------------------|-------------------|---------------------|-----------------|---------------|-------------------|---------------------|------------|---------------|-----|
|  | Eurostat data       | SILC data         |                     | Difference data | Eurostat data | HBS data          |                     | Difference |               |     |
|  |                     | simple statistics | taxes and transfers |                 |               | simple statistics | taxes and transfers |            |               |     |
|  | (1)                 | (2)               | CZ (3)              | HU (4)          | (5) = (4)-(3) | (6)               | HU (7)              | CZ (8)     | (9) = (9)-(8) |     |
| (a) full (15-74 years old)                           | 63.1                | 63.8              | 63.8                | 60.9            | -2.9          | 54.6              | 57.2                | 57.2       | 60.2          | 3.0 |
| (b) elementary school or less                        | 25.4                | 29.7              | 27.8                | 25.4            | -2.4          | 25.4              | 31.8                | 37.0       | 39.9          | 2.9 |
| (c) secondary education                              | 69.3                | 69.2              | 69.8                | 66.8            | -3.1          | 64.9              | 63.6                | 62.3       | 65.4          | 3.1 |
| (d) tertiary education                               | 80.1                | 79.2              | 78.4                | 75.5            | -3.0          | 75.9              | 75.1                | 71.3       | 74.1          | 2.8 |
| (e) elementary school or less, not full-time student |                     | 49.9              | 44.2                | 40.5            | -3.7          |                   | 40.9                | 46.8       | 50.4          | 3.6 |
| (f) secondary education, not full-time student       |                     | 72.4              | 72.7                | 69.6            | -3.1          |                   | 69.4                | 67.2       | 70.5          | 3.3 |
| (g) tertiary education, not full-time student        |                     | 82.7              | 82.0                | 79.1            | -2.8          |                   | 76.2                | 72.2       | 75.0          | 2.8 |
| (h) prime age (25-54 years old)                      | 87.3                | 87.6              | 88.2                | 84.9            | -3.3          | 80.1              | 84.1                | 81.2       | 85.1          | 3.9 |
| (i) single men                                       |                     | 92.2              | 91.0                | 88.7            | -2.3          |                   | 87.9                | 83.8       | 86.7          | 2.9 |
| (j) single women                                     |                     | 80.8              | 83.1                | 78.6            | -4.5          |                   | 80.6                | 80.5       | 84.7          | 4.2 |
| (k) married men                                      |                     | 97.8              | 97.0                | 95.7            | -1.3          |                   | 92.4                | 87.9       | 90.0          | 2.1 |
| (l) married women                                    |                     | 79.0              | 81.1                | 76.0            | -5.1          |                   | 76.4                | 74.0       | 80.0          | 6.0 |
| (m) women at child-bearing age (25-49)               | 78.0                | 77.8              | 80.5                | 75.2            | -5.3          | 73.3              | 77.4                | 77.9       | 83.7          | 5.8 |
| (n) elder (>=50)                                     | 47.3                | 48.7              | 50.2                | 49.0            | -1.2          | 35.2              | 43.2                | 41.1       | 43.5          | 2.4 |

Notes: Participation rates in 2008 for the Czech Republic and Hungary in %; percentage points in columns (5) and (10). Results using the overall sample of individuals aged 15-74 in rows (a) to (g), for prime-age individuals aged 25-54 in rows (h) to (m), and for elderly individuals in row (n). Eurostat statistics are reported in columns (1) and (6); the aggregate participation rates obtained from the datasets are in columns (2) and (7). Simulations are presented in columns (3) and (8) for the countries' own tax-benefit systems and in columns (4) and (9) for the other country's tax-benefit system. Columns (5) and (10) show how the labour force participation rate would change for the selected subgroups if the country adopted the other country's tax-benefit system.

**Table 3.5.: Population Shares (2008)**

|  | Full-time student (1) |      | Receives old-age pension (2) |      | Receives disability pension (3) |      | Highest level of education |      |      |      |      |      |
|--|-----------------------|------|------------------------------|------|---------------------------------|------|----------------------------|------|------|------|------|------|
|  | CZ                    | HU   | CZ                           | HU   | CZ                              | HU   | CZ                         | HU   | CZ   | HU   | CZ   | HU   |
| (a) full sample (15-74)                              | 12.3                  | 11.0 | 15.7                         | 19.1 | 5.3                             | 8.4  | 17.2                       | 26.2 | 68.8 | 57.2 | 14.0 | 16.6 |
| (b) elementary school or less                        | 40.8                  | 22.4 | 16.6                         | 23.1 | 6.9                             | 12.1 |                            |      |      |      |      |      |
| (c) secondary education                              | 6.0                   | 8.5  | 16.5                         | 16.8 | 5.5                             | 8.4  |                            |      |      |      |      |      |
| (d) tertiary education                               | 8.3                   | 1.5  | 10.7                         | 20.9 | 2.3                             | 2.8  |                            |      |      |      |      |      |
| (e) elementary school or less, not full-time student |                       |      | 28.0                         | 29.8 | 11.6                            | 15.6 |                            |      |      |      |      |      |
| (f) secondary education, not full-time student       |                       |      | 17.6                         | 18.4 | 5.9                             | 9.2  |                            |      |      |      |      |      |
| (g) tertiary education, not full-time student        |                       |      | 11.7                         | 21.2 | 2.3                             | 2.8  |                            |      |      |      |      |      |
| (h) prime age (25-54)                                | 3.1                   | 1.0  | 0.1                          | 1.0  | 5.1                             | 6.5  | 7.7                        | 16.2 | 74.0 | 63.6 | 18.3 | 20.2 |
| (i) single men                                       | 4.5                   | 2.6  | 0.0                          | 0.7  | 6.4                             | 6.8  | 8.5                        | 17.1 | 77.7 | 67.6 | 13.8 | 15.3 |
| (j) single women                                     | 6.7                   | 1.9  | 0.0                          | 0.4  | 6.2                             | 7.0  | 11.4                       | 18.0 | 68.0 | 55.7 | 20.7 | 26.3 |
| (k) married men                                      | 1.1                   | 0.0  | 0.1                          | 2.2  | 3.5                             | 5.8  | 4.6                        | 12.6 | 75.5 | 70.3 | 19.9 | 17.1 |
| (l) married women                                    | 1.9                   | 0.1  | 0.1                          | 0.4  | 4.9                             | 6.5  | 8.2                        | 17.7 | 73.5 | 59.7 | 18.3 | 22.6 |
| (m) women at child-bearing age (25-49)               | 4.3                   | 1.0  | 0.0                          | 0.1  | 3.8                             | 3.6  | 7.4                        | 16.1 | 72.6 | 58.7 | 20.0 | 25.2 |
| (n) elder (>=50)                                     | 0.1                   | 0.0  | 41.5                         | 46.0 | 9.2                             | 16.6 | 16.6                       | 28.0 | 71.0 | 54.6 | 12.4 | 17.4 |

Notes: The share of full-time students (1), individuals receiving an old-age pension (2) and individuals receiving a disability pension (3) and population shares by the highest level of education (4-6) as a percentage of the total 15-74 years old population (a) and as a percentage of the total number of individuals in selected subgroups (b-n).



### 3.5. Conclusion

Taking the example of two countries, the Czech Republic and Hungary, this paper investigates the extent to which cross-country differences in aggregate participation rates can be explained by the disparity in their tax-benefit systems. We first replicate for Czech household-level data the labour supply estimation for Hungary presented in Benczúr et al. (2014) and use the two perfectly comparable estimates of labour supply at the extensive margin to simulate how the aggregate participation rate would change in one country if it adopted the other country's tax and social welfare system.

Our estimation results yield similar labour supply elasticities for both countries, suggesting that individual preferences are essentially identical in the two countries analysed. Consistently with previous findings, lower educated individuals, the elderly and married women (or women of child-bearing age) are the most responsive to tax and transfer changes.

The simulation results show that about three percentage points out of the total difference of 6.6 percentage points in the participation rates of the 15–74 years old population can be explained by differences in the tax-benefit systems. The simulated effects are quasi-symmetric for almost all subgroups, meaning that if the Czech system was adopted, the Hungarian participation rate of a specific subgroup would increase by about the same number of percentage points as the Czech participation would decrease by if the Hungarian system were implemented. The largest difference explained by the difference in the tax-benefit systems is identified for married women and women of childbearing age. This is related to the more generous maternity benefit system in place in Hungary as compared to the Czech Republic. The “unexplained” part of the difference is also likely to be affected by some elements of the transfer system that we do not explicitly control for, most importantly the differences in the old-age and disability pension schemes.

Obviously, these results cannot be directly generalised to other countries. First, it is possible that individual preferences differ more across countries with different cultural and historical backgrounds or institutional structures even within narrowly defined sub-populations. Second, cross-country differences in the composition of the working-age population may lead to different results even following a similar shock to net income levels. Nevertheless, the exercise presented in this paper sheds some light on how important the tax-benefit system can be in explaining the differences in labour force participation between two otherwise similar countries.

Finally, it is important to note that we do not suggest that Hungary should blindly adopt the Czech tax and transfer system in order to increase labour force participation. First, even though the Czech participation rate is close to the EU15 average (the usual benchmark for catching-up EU member states), the Czech system might not be optimal in every respect. Second, governments' redistributive preferences may be different and therefore the optimal policy may also differ. The Czech

### 3.5 Conclusion

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system is still a good benchmark for Hungary, as it provides a realistically achievable alternative for Hungarian policymakers.





# A. Appendix

## A.1. Appendix to Chapter 1: Descriptive statistics

**Table A.1.:** Descriptive statistics of the variables

| Definition of the variable                          | 2005-2008 |       |      | 2009-2011 |       |      |
|---|-----------|-------|------|-----------|-------|------|
|   | N         | mean  | std. | N         | mean  | std. |
| short contract dummy (duration is one year or less) | 309 077   | 0.26  | 0.44 | 186018    | 0.29  | 0.46 |
| loan size / total assets (in logs)                  | 206 095   | 4.77  | 1.70 | 125 648   | 4.86  | 1.62 |
| firm level variables                                |           |       |      |           |       |      |
| employment (in logs)                                | 287 161   | 2.31  | 1.35 | 172 872   | 2.30  | 1.35 |
| foreign ownership dummy (over 50%)                  | 287 138   | 0.06  | 0.24 | 172 864   | 0.06  | 0.23 |
| importer dummy                                      | 309 077   | 0.21  | 0.41 | 186 018   | 0.19  | 0.39 |
| exports share in sales                              | 201 884   | 0.07  | 0.20 | 122 974   | 0.07  | 0.20 |
| firms' real capital (in logs)                       | 206 180   | 0.36  | 0.25 | 125 661   | 0.36  | 0.25 |
| current assets / total assets                       | 206 180   | 0.62  | 0.26 | 125 661   | 0.62  | 0.27 |
| total assets (in logs)                              | 206 180   | 11.34 | 1.88 | 125 661   | 11.46 | 1.89 |
| profits / total assets                              | 206 180   | 0.02  | 0.36 | 125 661   | 0.02  | 0.32 |
| bank level variables                                |           |       |      |           |       |      |
| foreign ownership dummy (over 50%)                  | 250 217   | 0.95  | 0.21 | 152 320   | 0.94  | 0.23 |
| total assets (in logs)                              | 250 217   | 14.15 | 0.96 | 152 320   | 14.49 | 1.01 |
| bank capital ratio                                  | 250 217   | 0.09  | 0.04 | 152 320   | 0.09  | 0.04 |
| bank liquidity ratio                                | 250 217   | 0.13  | 0.06 | 152 320   | 0.17  | 0.09 |
| profits / total assets                              | 250 217   | 0.01  | 0.02 | 152 320   | 0.00  | 0.01 |
| doubtful loan ratio                                 | 250 014   | 0.57  | 0.04 | 152 202   | 0.61  | 0.04 |

*Notes: The Table provides descriptive statistics on the variables of the dataset used in the estimations. The statistics, – number of observations, mean and standard deviation – are given for two periods, 2005-2008 and 2009-2011. The table groups variables related to the firms and related to the banks providing the loans separately.*

## A.2. Appendix to Chapter 2: Summary of cash transfers and tax expenditures taken into account in the estimation

This Appendix summarizes the basic features of tax expenditures and the cash transfers and tax expenditures taken into account in the estimation. In particular, we discuss child care (family) benefits and unemployment (welfare) benefits. We treated old-age and disability benefits as exogenous and, accordingly, did not include these benefits in the summary. This rests on the assumption that if an individual is entitled for these benefits (due to age or health status), we will observe that he/she is a recipient. This looks like a natural assumption in the case of disability benefits. In the case of old-age benefits, this treatment is justified by the fact that during the sample period old-age pension recipients were allowed to work without any penalty. Thus they did not face a choice between pensions and earnings.

### 1. Tax expenditures in the PIT

- a) *Employee tax credit (adózátrás)*,<sup>1</sup> ETC is a non-refundable tax credit applying to wage income. The ETC was modest in size until its expansion in 2002. During the period 2003-2011 it made the minimum wage nearly PIT-free. The ETC was phased out in most years at a rate of 9% in an income range around the average wage. Until its abolishment in 2012, its exact parameters were adjusted each year.
- b) *Family tax credit (családi adókedvezmény)*. The Hungarian PIT has been an individual-based (as opposed to a family-based) tax system during the sample period. One of the parents can deduct the family tax credit from his or her tax payment (or both can share the credit) based on the number of children in the household. Starting in 2006, families with one or two children were not eligible for the tax credit (until the tax credit was expanded in 2011).
- c) *Other tax credits* were abundant in the tax code until 2006; since then they have been gradually eliminated. We use information in the Household Budget Survey to assess the tax credits each individual can take advantage of.
- d) *Tax base issues*. During the sample period, insurance-based benefits were generally treated as wage income by the tax code while universal benefits were tax exempt. During the years 2007-2010 pension income constituted part of the tax base although it was

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<sup>1</sup>There is considerable heterogeneity in the official and scientific publications regarding the English translation of the various benefits. In this table we chose to use the simplest English translations that reflect the nature of the given benefit; we included the official Hungarian designations so that the benefits can easily be identified.

not taxed itself (it pushed other incomes into the upper tax bracket). Benefits 2c and 2d were treated similarly during the whole sample period.

## 2. Family benefits

- a) *Maternity benefit (TGYÁS)* is an insurance-based benefit that mothers are entitled to receive for 5 months around child-birth. Its condition is current employment (at the time of applying for the benefit). The monthly benefit is equal to 70% of past monthly wage. The recipient may not engage in paid work while receiving this benefit. No couple can receive two of benefits 2a-d at the same time.
- b) *Child-care benefit I (GYED)* is an insurance-based benefit that one of the parents is entitled to receive until the second birthday of the youngest child. Its condition is at least 12 months of employment in the 24 months before the child is born. The monthly benefit is equal to 70% of past monthly wage but it may not exceed 140% of the minimum wage. The recipient may not engage in paid work while receiving this benefit. No couple can receive two of benefits 2a-d at the same time.
- c) *Child-care benefit II (GYES)* is not conditional on employment (social insurance) history. One of the parents is entitled to receive the benefit until the third birthday of the youngest child. The benefit is pegged to the so-called ‘minimum pension benefit’, equal to HUF 28500 (around 40% of the minimum wage) in 2008. Recipients are restricted from working full time in the first year of this benefit. (The employment restrictions were loosened for the second and third year during the period of study.) No couple can receive two of benefits 2a-d at the same time.
- d) *Child-care benefit III (GYET)*: A parent is entitled to this benefit if he or she raises at least 3 children until the 8th birthday of the youngest child, independently of employment (social insurance) history. The benefit is pegged to the ‘minimum pension benefit’ (see 2c). Recipients of this benefit are restricted from working full time. No couple can receive two of benefits 2a-d at the same time.
- e) *Family supplement (sometimes called ‘family allowance’; családi pótlék)* is a universal benefit all families with children are entitled to receive. The sum of the benefit depends on the number of children, whether there are twins among the children, and whether any of the children is chronically ill. It was equal to HUF 12,200 (around 18% of the minimum wage) for a family with one child in 2008.

## 3. Unemployment benefits

- a) *Unemployment benefit I (1998-2005: munkanélküli járadék; 2006-: álláskeresési járadék):* Individuals who lost their jobs are eligible for the insurance-based unemployment benefit (renamed as ‘job-seekers’ benefit’ in 2006). Its maximum duration was shortened from 12 months to 9 months in 2000. Until 2006 it was equal to 65% of the previous wage (capped at 180% of the ‘minimum pension benefit’, see 2c). After 2006 it had two phases. The first phase lasted 3 months, during which the recipient received 60% of his/her past wage (capped at 120% of the minimum wage). The second phase lasted 6 months, during which the benefit was equal to 60% of the minimum wage. (If the individual did not have a full employment history in the four years before the job loss, the duration of the benefit could be shorter. The second phase was abolished in 2012.)
- b) *Unemployment benefit II (2003-2005: álláskeresést ösztönző juttatás; 2006-: álláskeresési segély):* Established in 2003, this was a fixed-sum benefit for individuals whose unemployment benefit I expired but still did not find a job. It was conditional on cooperation with the local unemployment administration. Between 2003-2005 the benefit lasted a maximum of 6 months; it was reduced to 3 months in 2006. From that year onwards the benefit was equal to 40% of the minimum wage. (It was abolished in 2012.)
- c) *Pre-retirement unemployment benefit (Nyugdíj előtti álláskeresési segély):* Individuals are entitled for this insurance-based benefit (which used to be a sub-case of benefit 3b after 2006) if they lose their job in the five years before the statutory pension age. The benefit is equal to 40% of the minimum wage. The benefit payment is suspended if the individual finds employment.
- d) *Regular social benefit (1998-2000: jövedelempótló támogatás; 2001-: rendszeres szociális segély)* is a welfare benefit individuals can receive if they are not eligible to any other unemployment (or disability or child-care) benefit (any more). For most of the sample period it was means-tested. The details of the means-testing changed in 2006. After 2006 the benefit supplemented a family’s income to 90% of the ‘minimum pension benefit’ per consumption unit but could not exceed the net minimum wage. (Its predecessor in the years 1998-2000 was a fixed-sum transfer and it was succeeded by a fixed-sum transfer in 2010.)



## A.3. Appendix to Chapter 3: Description of the Welfare Benefit Systems

### A.3.1. Hungary

*Maternity benefit* is paid to insured women who are away from work as a result of pregnancy. In Hungary, it is an insurance-based benefit, conditional on past employment. It is divided into three or four parts. The first part (“*Maternity benefit*”), equal to 70% of the past monthly wage, is received by mothers for 5 months around childbirth. The second part (“*Child-care benefit I*”) is another insurance-based benefit of the same amount that one of the parents collects until the second birthday of the youngest child. The third part (“*Child-care benefit II*”) is not conditional on employment or insurance history. One of the parents is eligible for the benefit until the third birthday of the youngest child. The fourth part (“*Child-care benefit III*”) is available to a parent if he or she raises three children until the eighth birthday of the youngest child, independently of employment or insurance history. These benefits are mutually exclusive and recipients are restricted from working full time.

*Unemployment benefits (UB)* are also insurance-based benefits. Depending on unemployment spell and eligibility, two types of benefits exist: *UB I* has a maximum duration of 9 months, split into two phases. In 2008, the recipient received 60% of the previous wage during the first half of the period (not longer than 3 months) and 60% of the minimum wage during the remaining phase; *UB II* is a fixed-sum benefit for individuals whose *UB I* expired without finding a job or for those who are not eligible for *UB I*. It is conditional on cooperation with *the* local unemployment administration. The benefit is equal to 40% of the minimum wage. Individuals who lose their job in the period of 5 years prior to the statutory pension age are entitled to pre-retirement unemployment benefit. This insurance-based benefit is equal to 40% of the minimum wage and is suspended if the individual finds a job.

*Family allowance* is a universal benefit for families with children. The amount of the benefit depends on the number of children. The benefit is not conditional on work.

*Regular social benefit* is a welfare benefit that an individual may receive if he or she is not eligible for any other unemployment, disability or child-care benefits. After 2006 the benefit supplemented the family income to 90% of the minimum pension benefit per consumption unit, but the total could not exceed the net minimum wage.

### A.3.2. Czech Republic

In the Czech Republic, the social security system consists of social insurance benefits, state social support and social assistance. Benefit is not subject to income taxation.

The social insurance benefits consist of unemployment, sickness and pension insurance benefits. *Maternity benefit* is part of the sickness contribution provided to mothers until 6 months after childbirth. The amount of the benefit is equal to 69% of the previous gross income.

Unemployed persons may receive *unemployment benefits* for a maximum period of 6 months. The amount of the benefit is 50% of the net previous work income in the first 3 months and 45% in the following months of the support period, while a maximum amount is applied. The support period is longer for older job seekers.

The state social support is targeted mainly at families with children and is based on the concept of the living minimum as stipulated by legislation. The income of the whole family is tested for the purposes of the *child allowance* (family benefit, “*prídavek na dítě*”), the *social allowance* (“*socialní příplatek*”) and the *housing allowance* (“*příplatek na bydlení*”). Income is not examined for the *parental allowance*, the foster care benefit, the birth grant and the funeral grant. The parental allowance is usually provided to mothers caring daily for a child until the age of 2, 3 or 4, but gainful activity is allowed.

Social assistance (*assistance in material need*) is the last resort for persons or families who are not able to afford the minimum living requirements as recognised by the state in the so-called living minimum (i.e. those in material need). The amounts of the living minimum are different for single persons, for first and other adult household members and for children. Social assistance gives preferential treatment to those recipients of benefits in material need who are actively seeking work or are working. For persons who are not eligible for material need assistance, the subsistence minimum is defined to cover personal needs on a survival level. Housing needs are covered separately by the *housing allowance* in state social support and by a *housing supplement*, which is a part of social assistance.



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## **Essais sur deux enjeux majeurs des pays d'Europe de l'Est: l'endettement en devises étrangères et l'offre de travail**

Cette thèse traite deux sujets distincts, les deux représentant des enjeux importants pour un grand nombre de Pays d'Europe Centrale et Orientale (PECO). La première partie porte sur les emprunts en devises étrangères. Plusieurs études antérieures montrent que dans de nombreux PECO, l'endettement en devises étrangères a augmenté de manière considérable avant la crise et est devenu un enjeu majeur pour les entreprises, les ménages et pour la politique budgétaire et monétaire. Pour évaluer les risques associés à l'endettement excessif en devises étrangères, nous étudions la volonté des entreprises d'apparier la composition en devises de leurs actifs et leurs passifs ainsi que leurs incitations à dévier de l'appariement parfait. Nos résultats fournissent des preuves solides à l'appui du rôle de la couverture naturelle. Néanmoins, ce dernier n'est pas le motif principal d'endettement en devise étrangère: le motif de couverture naturelle n'explique qu'environ 10 à 20 pour cent de la dette totale en devises étrangères des entreprises avant et pendant la crise, respectivement. La plus grande partie de la dette en devises étrangères correspondrait, au moins en Hongrie, à des positions de carry trade détenues par des sociétés non financières.

La deuxième partie de la thèse est consacrée à l'exploration des liens entre les systèmes socio-fiscaux et l'offre de travail à la marge extensive. Le deuxième chapitre propose une nouvelle stratégie de modélisation de l'offre de travail comme alternative aux deux approches dominantes basées sur le calcul marginal et les modèles d'utilité aléatoire. Finalement, le dernier chapitre utilise ce modèle pour quantifier la part de la différence entre les taux d'activité tchèque et hongrois qui peut être expliquée par les divergences des systèmes d'imposition et de protection sociale. Les estimations donnent des élasticités d'offre de travail similaires, ce qui suggère que les préférences individuelles sont essentiellement identiques dans les deux pays. Nos résultats montrent que la moitié de l'écart entre les taux d'activité s'explique par les différences des systèmes socio-fiscaux.

*Mots-clés:* décisions d'emprunt; asymétrie de devises; carry trade; crise financière; offre de travail, système fiscal-social; microsimulation.

## **Essays on two central issues in Central and Eastern European countries: foreign currency indebtedness and labour supply**

This thesis deals with two distinct topics, both of them representing central issues for many Central and Eastern European (CEE) countries. The first part of the thesis focuses on foreign currency (FX) lending. Several previous studies point out that in many CEE countries, FX borrowing rose significantly before the crisis and has become a major challenge for firms, households and for fiscal and monetary policy. To evaluate the risks associated with excessive FX indebtedness, we investigate firms' willingness to match the currency composition of their assets and liabilities and their incentives to deviate from perfect matching. Our results provide strong evidence to support the role of natural hedging, however, it is not the primary motivation for firms to choose foreign currency: it explains only about 10 percent of the overall corporate FX debt during the pre-crisis and 20 percent during the post-crisis periods. Most likely, the largest part of the corporate FX debt, at least in Hungary, corresponds to open carry trade positions held by non-financial corporations.

The second part of the thesis is devoted to exploring the links between tax-benefit systems and labour supply at the extensive margin. The second chapter presents an alternative modelling strategy of labour supply to the two dominating approaches based on marginal calculus and on random utility models. Finally, the last chapter uses this model to quantify the difference between the Hungarian and the Czech participation rates that can be attributed to differences in taxation and welfare benefits. We find that the estimated labour supply elasticities for the Czech Republic are very close to the results for Hungary, suggesting that, at least in this dimension, individual preferences are similar in the two countries. Results suggest that about one-half of the total difference in the participation rates can be explained by differences in the tax-benefit systems.

*Keywords:* borrowing decisions; currency mismatch; carry trade; financial crisis; labour supply, tax-benefit systems; microsimulation.