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Résumé

Cet article estime l'effet de l'impôt sur les revenus et des transferts sociaux sur l'offre de travail à la marge extensive, i.e. la participation au marché du travail. Nous étendons les méthodes structurelles existantes en tenant compte simultanément de l'effet des impôts et des transferts. Les revenus non salariaux incluent le montant (hypothétique) des transferts que l'individu receverait sans travailler. Quant au salaire, il est remplacé par la différence entre le montant du salaire net et le montant des transferts perdus en devenant salarié (*"gains to work", GTW*) . Afin d'intégrer ces composantes dans la contrainte budgétaire, nous avons recours à un modèle impôts-prestations détaillé. En utilisant les données de *l'enquête sur le budget des ménages hongrois (HKF)*, nous trouvons que la probabilité de participation est fortement influencée par les transferts et le GTW, en particulier pour les individus peu qualifiés et les personnes âgées. De plus, le même changement dans le salaire net conduit à un changement beaucoup plus important dans le GTW pour les bas salaires, ce qui les rend encore plus sensibles aux impôts. Dans l'ensemble, nous constatons qu'une seule équation peut expliquer la grande partie de l'hétérogénéité des individus en termes de sensibilité aux impôts et aux transferts. Nos estimations paramétriques peuvent être aisément utilisées pour étudier les effets des réformes du système d'imposition et des transferts sur le bien-être et le comportement des individus.

Mots-clés: décision de participation, imposition, transferts sociaux

Codes JEL: H24, H31, H53, I38, J21

Abstract

This paper estimates the effect of income taxation and transfers on labour supply at the extensive margin, i.e., the labour force participation. We extend existing structural form methodologies by considering the effect of both taxes and transfers. Non-labour income contains the (hypothetical) transfer amount someone gets when out of work, while the wage is replaced by the difference between net wages and the amount of lost transfers due to taking up a job (gains to work). To incorporate these components of the budget set, we employ a detailed tax-benefit model. Using data from the *Hungarian Household Budget Survey (HKF)*, we find that participation probabilities are strongly influenced by transfers and the gains to work, particularly for low-skill groups and the elderly. Moreover, the same change in the net wage leads to a much larger change in the gains to work for low earners, making them even more responsive to wages and taxation. Overall, we find that a single equation can capture a large heterogeneity of individual responsiveness to taxes and transfers. Our parametric estimates can be readily utilized in welfare evaluations, or microsimulation analyses of tax and transfer reforms.

Keywords: participation decision, taxation, transfers

JEL classification: H24, H31, H53, I38, J21

Non-technical summary

The recent slump of European economies has highlighted the importance of labour market incentives: there is both a general need to boost labour market activity and a specific concern about the negative impact of ”fiscal austerity” on employment. A key ingredient to understand and predict the impact of such reforms is the response of labour supply.

This paper presents a unified structural approach to estimate the impact of taxes and transfers on the participation decision (the extensive margin of labour supply). Using individual (household-level) data and a multitude of tax and transfer reforms, we are able to control for shifts in labour demand and focus squarely on the determinants of labour supply. By utilizing a detailed tax-benefit (microsimulation) model, we are able to describe the budget set of individuals accurately. Moreover, our approach allows a general assessment of the efficiency and effectiveness of government interventions into the labour market, and more importantly, a micro-based prediction of the impact of tax and welfare reforms (as done in our related work in (Benczúr et al. (2011) and Benczúr et al. (2012)).

To incorporate the detailed impact of taxes and transfers on the budget set, we show that one has to modify the net wage by deducting the amount of lost transfers to get the measure which determines the participation decision (the gains to work). This leads to a modified but otherwise standard structural probit specification. Transfers are, however, not fully observed: for those who work, we do not observe the amount of transfers they would receive if they were to lose their job; while for those who do not work, we do not observe the amount of transfers they would lose if they were to take up a job. The paper of van Soest (1995) discusses the issue of the first aspect, but data limitations prohibit the author from its proper incorporation. We overcome both obstacles by using a detailed tax-benefit model based on individual characteristics and the welfare system's particulars for every given year. This allows the full and precise determination of both types of unobserved transfer changes.

Hungary provides a particularly interesting case for estimating the determinants of labour market activity. Participation rates are among the lowest in the EU. This has been often identified as a key bottleneck to real convergence (see e.g. Kátay (2009)). Motivated by this, Hungary has adopted sweeping changes in its tax and transfer system: in 2009-10, it increased substantially the upper limit of the lower tax bracket and cut employer contributions, then started to introduce a near flat personal income tax scheme starting in 2010 and to reduce social benefits starting in 2011. At the same time, the need for fiscal consolidation (being under the Excessive Deficit Procedure in 2004-2013) has put tax revenues and welfare benefits under pressure.

We carry out our estimation 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. 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.

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. This latter is due to the fact that the same change in the net wage leads to a very different change in the GTW if lost transfers are a different share of the net wage. 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. Our estimates imply an aggregate labour supply elasticity of 0.28, quite in line with the 0.25 'consensus estimate' of Chetty et al. (2012). 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%. This is a sum of a 2.09% decline due to the elimination of the employee tax credit, a 0.34% decrease due to higher social contributions and an offsetting increase of 1.51% due to tax rate cuts. Overall, this illustrates both the usefulness of our parametric approach for assessing the impact of tax and transfer reforms, and the economic significance of our parameter estimates.

1 Introduction

The recent slump of European economies has highlighted the importance of labour market incentives: there is both a general need to boost labour market activity and a specific concern about the negative impact of "fiscal austerity" on employment. A key ingredient to understand and predict the impact of such reforms is the response of labour supply. This paper presents a unified structural approach to estimate the impact of taxes and transfers on the static participation decision (the extensive margin of labour supply). Using individual (household-level) data and a multitude of tax and transfer reforms, we are able to control for shifts in labour demand and focus squarely on the determinants of labour supply. By utilizing a detailed tax-benefit (microsimulation) model, we are able to describe the budget set of individuals accurately. Moreover, our approach allows a general assessment of the efficiency and effectiveness of government interventions into the labour market, and more importantly, a micro-based prediction of the impact of tax and welfare reforms.

There are a number of existing studies which establish that taxes and the welfare system influence the participation decision, using either a static or a dynamic approach. While the latter can handle a multitude of important issues better (like the marriage and fertility decision, human capital accumulation and state dependence in tastes for work), it is computationally challenging to estimate and simulate such models on large micro data.[1] For this reason, we concentrate on static estimates. The literature exhibits a notable heterogeneity in terms of implied elasticity measures.[2] Arrufat and Zabalza (1986) do a cross section estimation on the U.K. General Household Survey dataset, and find a participation elasticity (the change in the probability of being active in response of a unitary shock in net wages) of 1.41 for married women. Dickert et al. (1995), conducting a cross-section estimation on the Survey of Income and Program Participation (SIPP) to analyse a large expansion of the Earned Income Tax Credit (EITC) in the U.S., find an elasticity of $\eta = 0.2$ for single parents. Using Dutch data, van Soest (1995) obtains an elasticity of 0.08-0.15 for men and 0.47-1.03 for women. Eissa and Liebman (1996) follow a program evaluation methodology (difference in differences) using the Current Population Survey to analyse the same episode of EITC expansion. They find that single mothers increased their participation rate by 2.8 percentage-points relative to single women without children. Kimmel and Kniesner (1998) adopt a panel estimation on SIPP, and find elasticities of $[0.6; 2.4; 1.8; 1.1]$ for single men, single women, wives and husbands respectively. Finally, Aaberge et al. (1999) follow a cross section estimation based on the Survey on Household Income and Wealth (Italy), and obtain average elasticities for men and women as $[0.04; 0.65]$ respectively.

From the point of view of tax and transfer reform impact assessment, these findings have important shortcomings. First, most of them focus on special subgroups and tend to follow a reduced form approach (program evaluation methodology, see Moffitt (2002) for a review). Though such approaches are capable of precisely estimating the impact of a particular tax or transfer reform

[1] See Keane et al. (2011) for a recent survey on the dynamic discrete choice approach.

[2] The recent survey of Keane (2011) also documents a large heterogeneity in labour supply elasticities, both in static and dynamic frameworks. The results the paper reports for men, however, are mostly for the hours worked decision; while for women, they are usually the sum of the intensive and the extensive margin responses.

episode, they are not suitable for evaluating the impact of future (hypothetical) scenarios. There is also a substantial heterogeneity in the way after-tax wages are controlled for (if at all). Meyer and Rosenbaum (2001) is an example of a structural approach, but is not suitable for simulations either: wages are proxied, so the results do not imply a wage elasticity.

Second, the existing literature usually focuses on either taxes or transfers (van Soest (1995) is an important exception). Though the meta-analysis of Chetty et al. (2012) provides a "new consensus estimate" of (Hicksian) extensive margin elasticities of 0.25, this result still does not control for the entire tax and transfer system. As argued by Blundell (2012), it is important to take taxes and transfers into account simultaneously and combine them into effective 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.

We base our empirical approach on an extension of the standard labour supply model, which incorporates both the marginal and participation tax rate aspect of transfers. This builds on the observation that jobs usually have a fixed minimum size (half- or full-time), which implies that an interior solution at a too low number of hours is infeasible. In that case, the labour supply choice of individuals is determined by comparing their utility level in and out of the labour market.

The underlying theory - presented in Section 2 - leads to a structural probit 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. Notice that the gains to work is influenced both by taxes and (lost) transfers. Transfers are, however, not fully observed: for those who work, we do not observe the amount of transfers they would receive if they were to lose their job; while for those who do not work, we do not observe the amount of transfers they would lose if they were to take up a job. The paper of van Soest (1995) discusses the issue of the first aspect, but data limitations prohibit the author from its proper incorporation. We overcome both obstacles by using a detailed tax-benefit model, based on individual characteristics and the welfare system's details for every given year. This allows the full and precise determination of both types of unobserved transfers.

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 point lower than the EU28 average. By 2014 the gap narrowed to 6.8%, still yielding a position of 24/28.[3] This has been often identified as a key bottleneck to real convergence (see e.g. Kátay (2009)). Motivated by this, Hungary has adopted sweeping changes in its tax and transfer system: in 2009-10, it increased substantially the upper limit of the lower tax bracket and cut employer contributions, then started to introduce a near flat personal income tax scheme starting in 2010 and to reduce social benefits starting in 2011. At

[3] The source is Eurostat, population between 15 and 64.

the same time, the need for fiscal consolidation (being under the Excessive Deficit Procedure in 2004-2013) has put tax revenues and welfare benefits under pressure.

We carry out our estimation 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. The estimation process - described in Section 3 - 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. In Section 4, 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 5 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,[4] 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. (2012). 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%. This is a sum of a 2.09% decline due to the elimination of the employee tax credit, a 0.34% decrease due to higher social contributions and an offsetting increase of 1.51% due to tax rate cuts.[5] Overall, this illustrates both the usefulness of our parametric approach for assessing the impact of tax and transfer reforms, and the economic significance of our parameter estimates.

2 Theory

The usual approach is to define the reservation wage, which is the threshold for accepting a job offer. Let us start from a standard utility maximization problem:

$$\begin{cases} \max \ f\left(c, 1-l\right) \\ \text{s.t.: } c + w\left(1-l\right) = w + T, \end{cases}$$

where c is consumption, l is labour, w is the wage, and T denotes transfers and other non-labour income. The total time endowment is normalized to 1, so leisure is $1-l$. The optimality condition

[4] As argued for example by Kátay (2009), Hungary's labour participation deficit compared to other EU members is mostly due to these special groups.

[5] The sum of the effects of these measures may differ from the total effect due to interactions.

can be written as

$$f_{1-l}\left(c,1-l\right)=wf_{c}\left(c,1-l\right).$$

The reservation wage is the lowest wage that induces nonzero hours worked. It corresponds to the case where $1-l^{*}=1$, i.e., tangency occurs exactly at zero hours worked. Then $c=T$, so

$$w_{res}=\frac{f_{1-l}}{f_{c}}\left(T,1\right)$$

defines the reservation wage. The participation decision is then determined by $w\geq w_{res}$. Loglinearising the right hand side, or working with specific functional forms for f,[6] we get

$$\log w\geq\log\chi+\psi\log T. \tag{2.1}$$

To allow for individual heterogeneity, we assume that individuals differ in their utility of leisure term (χ) and then expand $\log\chi_{i}$ as $Z_{i}A'+\varepsilon_{i}$, where Z_{i} is a vector of observable individual characteristics and $\varepsilon_{i}\sim N\left(0,\sigma^{2}\right)$:

$$\log w_{i}-Z_{i}A'-\psi\log T_{i}\geq\varepsilon_{i}.$$

The probability of someone working given a wage offer w_{i}, non-labour income T_{i} and individual characteristics Z_{i} is then

$$P=\Phi\left(\frac{\log w_{i}-Z_{i}A'-\psi\log T_{i}}{\sigma}\right)=\Phi\left(\gamma\log w_{i}+Z_{i}\alpha'-\bar{\psi}\log T_{i}\right), \tag{2.2}$$

yielding the standard structural probit specification.

The next step is to add taxes and transfers. One the one hand, we have to modify the wage rate by the effective tax rate (marginal rate, at zero labour income), 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 ΔT 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 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 reservation wage is infinite (this can also be established formally by total differentiation).

Instead, we redefine the reservation wage by constraining the participation decision to a fixed "job size" l^{*} – in our empirical specification, it will be a full time job.[7] This is a popular assumption

[6] The standard separable CES form $\frac{c^{1-\psi}-1}{1-\psi}+\chi\frac{(1-l)^{1-\phi}-1}{1-\phi}$ yields exactly (2.1). The growth-consistent utility function of the form $\frac{(c\cdot exp(g(1-l)))^{1-\psi}-1}{1-\psi}$ also yields (2.1), with an extra constraint of $\psi=1$.

[7] Once working, an individual may decide to work more than $l*$. We assume, however, that it is not known in advance whether there would be opportunities for overtime or performance bonuses, so the activity decision is determined by

in the discrete choice approach of the labour supply literature (dating back to van Soest (1995)). The reservation wage is thus set by the following comparison:

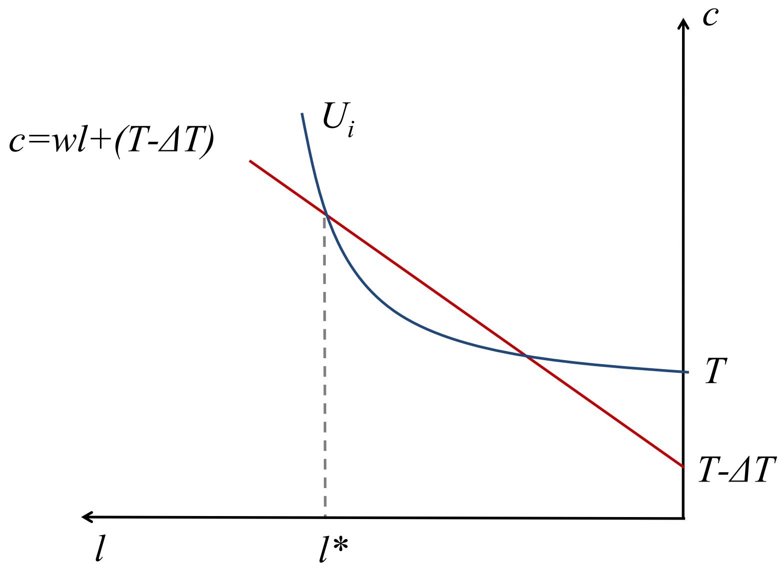
- Do not work: then $c = T, 1 - l = 1$, welfare is $f(T, 1)$.
- Work l^*: then $c = T - \Delta T + wl^*$, $1 - l = 1 - l^*$, welfare is $f(T - \Delta T + wl^*, 1 - l^*)$.

Introducing the notation $W = wl^* - \Delta T$ (gains to work, GTW), the comparison becomes:

$$f(T + W, 1 - l^*) \geq f(T, 1). \tag{2.3}$$

One can also give a simple graphical representation (see Figure 2.1): draw the indifference curve going through $(C = T, l = 0)$, find the point of this curve where $l = l^*$, and connect this with point $(C = T - \Delta T, l = 0)$. Its slope is then the reservation wage: at such a wage level, the individual is just indifferent between not working and getting the full amount of transfers ($C = T, l = 0$), or working l^* hours and getting only $T - \Delta T$ as transfers ($C = T - \Delta T + W, l = l^*$).

Figure 2.1: The reservation wage when there is a discrete drop in transfers



To derive a formal expression for the probability of being active (the analogue of (2.2)), let us linearise the left hand side of (2.3):

$$f(T + W, 1 - l^*) \approx f(T, 1 - l^*) + W f_c(T, 1 - l^*),$$

so the comparison becomes

$$W \geq \frac{f(T, 1) - f(T, 1 - l^*)}{f_c(T, l^*)}.$$

the base salary.

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

$$\log W - \psi \log T - \log \chi \geq \varepsilon,$$

yielding again a structural probit of the form

$$P = \Phi\left(\gamma \log W_i + Z_i \alpha' - \bar{\psi} \log T_i\right). \tag{2.4}$$

Let us compare the two structural probit equations (2.2) and (2.4). First, W_i in (2.4) represents the gains to work (from a full time job): $W_i = w_i l^* - \Delta T_i$, as opposed to the net wage w_i. Second, T_i is the hypothetical amount of transfers one gets (or would get) at zero hours worked.

From a practical point of view, T is not directly observable for the employed, since they get $T - \Delta T$; while ΔT is not observed for the inactive, since they get T. Using individual characteristics and the welfare system's details (for every given year), however, one can back up T and ΔT. This essentially requires a tax-benefit model (a microsimulation tool). For those who work, we determine T based on their characteristics and welfare regulations for the given year, and then obtain $\Delta T = T - T_{obs}$. For those who do not work, we determine ΔT by again applying welfare rules, while $T = T_{obs}$.

3 Econometric issues

We want to estimate a structural probit equation:

$$P\left(\text{employed/active}\right) = \Phi\left(\gamma \log W_i + Z_i \alpha' - \bar{\psi} \log T_i\right),$$

where $W_i = w_i l^* - \Delta T_i$. Here the vector Z_i contains individual characteristics which shift the labour supply of an individual. As usual in the literature on participation, there is a missing data issue: the wage is unavailable for those who do not work. The solution is to use a predicted W for the inactive: run

$$\log W_i = X_i \beta' + \mu_i$$

for the employed, and use the predicted wage $\hat{W} = X_i \widehat{\beta}$ for the unemployed. Here the vector X_i contains individual characteristics which are relevant for defining an individual's wage. 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 W error terms might be correlated. The solution is thus to adopt a Heckman-type correction, yielding a three step procedure.[8]

Given that our wage variable is gains to work, one can follow two alternatives. In variant A, the procedure is the following:

[8]Kimmel and Kniesner (1998) follows such an approach, for example.

1. Run a reduced form probit

$$P\,(\text{employed}) = \Phi\left(X_i\beta'_{RF} + Z_i\alpha'_{RF} - \psi_{RF}\log T_i\right).$$

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

$$\log W_i = X_i\beta' + \delta\lambda\left(X_i\widehat{\beta}'_{RF} + Z_i\widehat{\alpha}'_{RF} - \widehat{\psi}_{RF}\log T_i\right) + \mu_i.$$

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

$$P\,(\text{employed/active}) = \Phi\left(\gamma\widehat{\log W}_i + Z_i\alpha' - \bar{\psi}\log T_i\right).$$

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

$$P\,(\text{employed}) = \Phi\left(X_i\beta'_{RF} + Z_i\alpha'_{RF} - \psi_{RF}\log T_i\right).$$

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

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

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

$$P\,(\text{employed/active}) = \Phi\left(\gamma\widehat{\log W}_i + Z_i\alpha' - \bar{\psi}\log T_i\right).$$

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, her reported wage is less than the true annual wage. Alternatively, unreported wage income can also lead to a mismeasurement of wages. Notice, however, that we are in fact running an IV-probit in step 3, which offers a remedy to both of these problems (as long as there are variables in X_i which are excluded form Z_i, an issue we address in the data section).

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 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 the GTW ($\log W$) at least through the change in transfers term ΔT. For the structural probit equation (2.4) 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 γ 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 form 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.[9]

Finally, there is a tradeoff between adopting Variant A or B. The latter would seem more appealing, since it allows for $X \not\supseteq Z$, 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.[10] 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, is is much less of an issue in the structural probit equation.

Since our "wage" measure in the structural estimation is the GTW, the calculation of regular wage elasticities requires one more step. The structural probit gives us a log GTW coefficient γ. 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 GTW. This already implies some heterogeneity of elasticities. Even then, however, it is still the impact of a change in W, not w.

[9] 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).

[10] 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

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

so

$$\frac{\partial \Phi}{\partial \log w} = \frac{\partial \Phi}{\partial \log W}\frac{\partial \log W}{\partial \log w} = \frac{\partial \Phi}{\partial \log W}\frac{w}{w - \Delta T}. \tag{3.1}$$

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.

4 Data

We use data from the Hungarian Household Budget Survey (HKF), years 1998-2008. The dataset contains detailed income and consumption measures of broadly $25,000$ individuals per year in a repeated cross-section.[11]

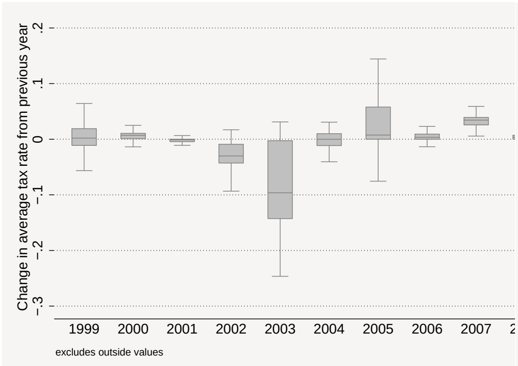
This period featured numerous policy measures on both income tax rates and transfers, providing us with sufficient exogenous cross-sectional and time-series variation in non-labour income and gains to work. Starting with taxes, Figure 4.1 shows how individuals' average tax rates would have changed if 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 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.

As for transfers, Figure 4.2 – adapted 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 4.2, 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.

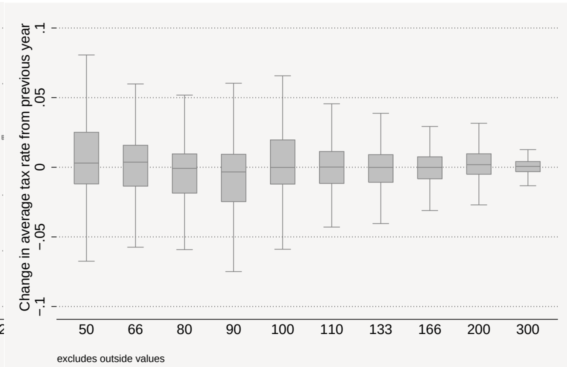
For our methodology to work, it is key to define the counterfactual transfers: First, how much would someone who is currently working receive in transfers if that individual is laid off? Second,

[11] 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.

Figure 4.1: 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.

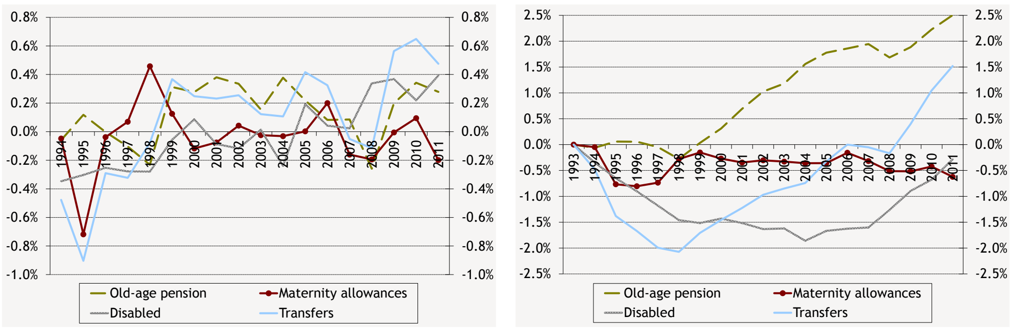
how much would someone who is currently inactive lose if that individual takes up a full time job? Calculating these measures requires the detailed coding of the full transfer system, basically a microsimulation model. We detail the major tax expenditure and cash transfer items in the Appendix. With one exception, the database contained all the relevant information to deduct the counterfactual transfer entitlements or losses of each individual. The exception was 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 main left hand side variable is labour force participation,[12] 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_{it} 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 (see

[12]It is the ”most typical” status for the given year, self-reported by survey respondents. Unemployment is defined along the ILO classification.

Figure 4.2: Decomposition of the aggregate participation rate*



(a) Year-to-year percentage point changes

(b) Cumulative changes

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

the first remark at the end of Section 3), and also as a source of additional variation to identify the parameter γ (remark two of the same section). 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 (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).

5 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).

5.1 Overall participation elasticities

Table 1 displays our baseline results, following the econometric methodology of Variant A. Panel A reports the estimates for the structural probit equation (2.4). Most point estimates have the expected sign and are significant. A higher GTW 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.

Table 1: Main results

	(A) Estimation results			
	participation (1)		employment (2)	
	coef.	*std. err.*	*coef.*	*std. err.*
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			
	dy/dx	*std. err.*	*dy/dx*	*std. err.*
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 then three years-old), full time student, education dummies (less then 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.*

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 GTW leads to a 2.9% increase in the probability of being active. As explained by equation (3.1), 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.

The conditional marginal effects presented in Table 1 are not directly comparable to the 'consensus' 0.25 value of aggregate net wage elasticity reported by Chetty et al. (2012): 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. (2012) indicating the percentage change in total employment to the same shock. To produce the equivalent of the exercise by Chetty et al. (2012), 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, probability 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.

5.2 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 presents two variants, a full and a restricted sample estimate. The full sample means that all observations are included (as in Table 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.

Notice that the net wage (or even the GTW) elasticity of activity is highly different across the three educational groups 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 γ and $\bar{\psi}$ in equation (2.4)).

If those two parameters are common across individuals, than labour supply elasticities at the intensive margin are also common: one can show that for a fixed income share $W/(W+T)$ and expenditure share $\alpha = c/(c + w\left(1-\bar{l}\right)$, the impact of a change in the net wage or transfers is the same on the hours worked decision of every individual. This homogeneity is however partial, since individuals with different gross wages (productivity) or transfers (non-labour income in general) will have different income and expenditure shares. When there is no non-labour income ($T = 0$), this homogeneity becomes even more complete, as the labour supply elasticity depends only on common parameters and original hours worked ($\alpha = \bar{l}$). So if individuals differ in their characteristics but their original hours worked is the same, so is their intensive margin labour supply elasticity. If utility is linear in consumption ($\psi = 0$), then the elasticity (of leisure) to net wages is common across all individuals (full homogeneity).

Table 2: Probit estimates and conditional marginal effects by subgroups

		full sample (1)		restricted sample (2)	
		dy/dx	*std. err.*	*dy/dx*	*std. err.*
elementary 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 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 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.*

Table 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 sizable 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 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.

Table 3: Probit estimates and conditional marginal effects by subgroups, prime-age subsample

		full sample (1)		restricted sample (2)	
		dy/dx	*std. err.*	*dy/dx*	*std. err.*
full prime-age 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, elementary 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, secondary 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, tertiary 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.*

Finally, Table 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.

Table 4: Conditional marginal effects by selected subgroups

		dy/dx	*std. err.*
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 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*

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 GTW. 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.

5.3 Economic significance of the estimated elasticities

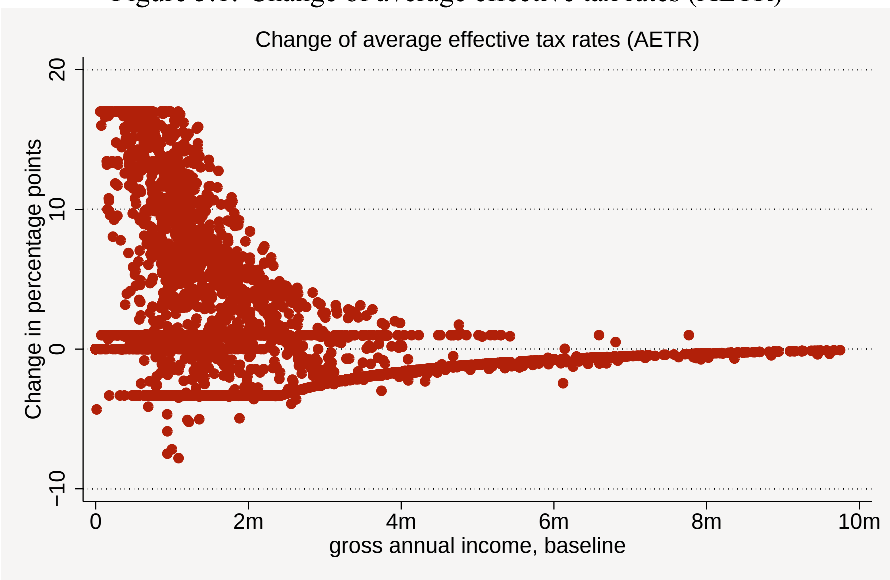
We now demonstrate how our results can be utilized for the simulation of the labour supply (participation) effect of a personal income tax and transfer reform. The main step is to calculate the probability of being active for a given hypothetical wage, tax and transfer system. First we obtain the pre- 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 the log of the GTW $(\log W)$ before and after the reform.

Equipped with these, we form

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

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 5.1: 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,[13] and a 1 percentage point increase in the social contribution rate. As illustrated by Figure 5.1, 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 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%, from which the elimination of the ETC is responsible for 2.09%,[14] the increase in social contributions leads to another 0.34% reduction, which are partly offset by an increase of 1.51% due to the rate cut.[15] Overall, this illustrates both the usefulness of our parametric approach for assessing the impact of tax and transfer reforms, and the economic significance of our parameter estimates.

6 Conclusion

This paper presents a first (at least to our knowledge) structural form estimation of labour supply at the extensive margin that incorporates the detailed impact of taxes and transfers on the budget set. We show that one has to modify the net wage by deducting the amount of lost transfers to get the measure which determines the participation decision (the gains to work). This implies, however, that the same change in the net wage leads to a very different change in the GTW if lost transfers are a different share of the net wage.

A major data obstacle is the need to determine counterfactual transfers: for those who work, we need the amount of transfers they would receive if they were to lose their job; while for those who do not work, we need the amount of transfers they would lose if they were to take up a job. Using a detailed tax-benefit model, we can obtain a full and precise determination of both types of unobserved transfers.

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 GTW. The most responsive subgroups are exactly the ones who are mostly responsible for Hungary's low participation rate (low-skilled, women at 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 demonstrated how our model can be utilized to calculate the labour supply shift of a complex personal income tax reform. In

[13] The average exchange rate in 2012 was 290 HUF per euro.

[14] 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.

[15] The sum of the effects of these measures may differ to the total effect due to interactions.

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 (2011)), and then it is embedded in a small general equilibrium macro model. With such a fully fledged model, we were able to evaluate at depth the 2011-12 Hungarian tax and transfer reforms as well (Benczúr et al. (2011) and Benczúr et al. (2012)).

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Appendix: 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 (adójóváírás),*[16] *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 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

[16] 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.

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 *i*nsurance-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 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.)

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