Income Taxation, Transfers and Labour Supply at the Extensive Margin☆

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Abstract

This paper estimates the effect of income taxation 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 sum of net wages and the amount of lost transfers due to taking up a job (gains to work, GTW). Using data from the *Hungarian Household Budget Survey (HKF)*, we find that participation probabilities are strongly influenced by transfers and the GTW, particularly for low-income groups and the elderly. Moreover, the same change in the net wage leads to a much larger change in the GTW for low earners, making them even more responsive to wages and taxation. 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

1. Introduction

This paper presents a unified parametric approach to estimate the impact of taxes and transfers on the participation decision (the extensive margin of labour supply). In our framework, participation probabilities are determined by the comparison of disposable income in

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and out of the labour force, consisting of the (often non-observed) amount of transfers and non-labour income an individual gets if not working and the gains to work (GTW; change in disposable income if accepting a job offer, the sum of net wages and lost transfers). Identification is achieved by utilizing a multitude of tax and transfer reforms. Unlike in the existing literature, our results allow 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 is a multitude of existing studies which establish that taxes and the welfare system influence the participation decision. There is, however, a notable heterogeneity in terms of implied elasticity measures. 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. 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.

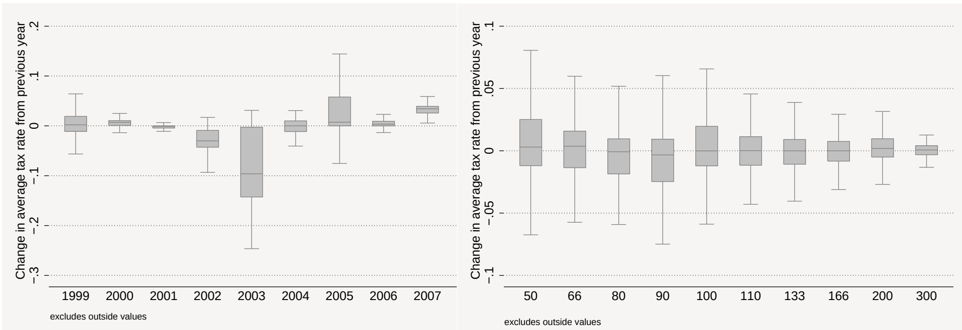
From our point of view, these findings have important shortcomings. First, most of them focus on special subgroups and tend to follow a reduced form approach (program evaluation methodology, see Moffitt (2002) for a review). Though such approaches are capable of precisely estimating the impact of a particular tax or transfer reform 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. Though the meta-analysis of Chetty et al. (2012) provides a "new consensus estimate" of extensive margin elasticities of 0.25, this result still does not necessarily 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 both marginal and average 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 an average tax rate: the total amount of lost transfers decreases the payoff from work, just like the average tax rate does.

One major reason for the lack of structural studies is that it is not obvious how to incorporate all the relevant features of the tax and transfer system into a theory-based framework of labour supply. This paper presents an extension of the standard labour supply model that can incorporate both the marginal and participation tax rate aspect of transfers, but at the expense of constraining the participation decision to a fixed job size. Jobs usually have a fixed minimum size (half-time, or in some cases even full-time), which implies that an interior solution at a too low number of hours might also be practically infeasible. In that case, the labour supply choice of individuals is determined by the average tax rate at her initial gross monthly earnings and the total amount of transfers. The overall summary measure in this case is the gains to work, which consists of the net wage (for the fixed size of the job) minus the amount of lost transfers.

We carry out our estimation on the Hungarian Household Budget Survey (HKF), containing detailed income and consumption measures of individuals for the years 1998-2008. Numerous policy measures on both income tax rates and transfers adopted during this period provide enough cross-sectional and time variation for the estimation of the elasticity of participation probabilities with respect to gains to work. Figure 1 show how individuals' average tax rates would have changed if their real income remained unchanged over time. It is seen that minor income tax changes occurred every year and major changed occurred in 1999 and between 2002 and 2005. The right graph show that tax changes affected lower income earners to a greater extent. As for the transfers, Figure 2 illustrates the impact of various transfer reforms on the Hungarian participation rate. The simple decomposition exercise of Kátay and Nobilis (2009) clearly demonstrates that transfer changes do impact the participation rate, providing us with sufficient exogenous variation in transfers to identify our specification.

Figure 1: Variation in the changes in average tax rates (atr)*

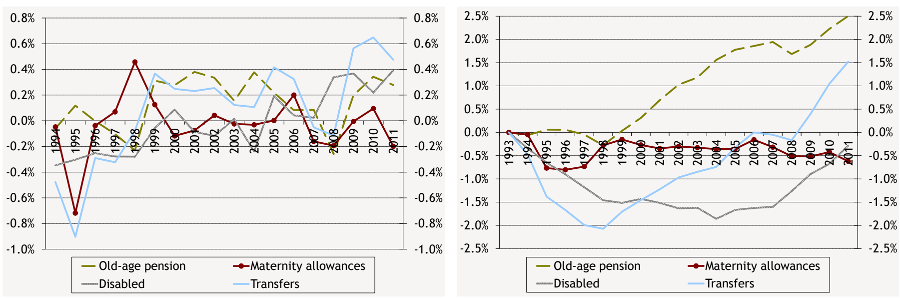


(a) yearly changes in atr (b) changes in atr by income categories (% of average income)

** 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.

Figure 2: Decomposition of the aggregate participation rate*



(a) year-to-year percentage point changes (b) cumulative changes

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

The underlying theory - presented in Section 2 - leads to a structural probit equation which relates participation probabilities to gains to work from a full time job, the total amount of non-labour income (including the hypothetical amount of transfers one gets or would get at zero hours worked) and other individual characteristics. The unobserved hypothetical amount of transfers are backed up using individual characteristics and the welfare system's details for every given year. 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 careful 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 impacts activity indirectly. In Section 4, we argue that county dummies and (once we control for individuals' lifecycle position with a large set of dummy variables) individuals' age are such variables.

Section 5 presents the estimation results. We find that a single equation can already explain a large heterogeneity of individual responsiveness to taxes and transfers: there are large differences among subgroups, driven partly by a composition effect, and partly by a different share of lost transfers in the GTW. The most responsive subgroups are low-skilled, (married) women at child-bearing age and elders, while prime-age higher educated individuals are practically unresponsive to tax and transfer changes at the extensive margin. As argued for example by Kátay (2009), Hungary's labour participation deficit compared to other EU members is mostly due to these special groups.

2. Theory

2.1. The underlying 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:

$$\begin{cases} \text{max } \frac{c^{1-\psi}-1}{1-\psi} + \chi\frac{(1-l)^{1-\phi}-1}{1-\phi} \\ \text{s.t.: } c + w\,(1-l) = w + T, \end{cases}$$

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 can be written as

$$\chi\,(1-l)^{-\phi} = wc^{-\psi}.$$

The reservation wage corresponds to the case where $1 - l^* = 1$. Then $c = T$, so

$$\chi = w_{res}T^{-\psi}$$

defines the reservation wage. The participation decision is then determined by $w \geq w_{res}$, or in logs:

$$\log w \geq \log\chi + \psi\log T.$$

Finally, we 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

$$P = \Phi\left(\frac{\log w_i - Z_i A' - \psi \log T_i}{\sigma}\right) = \Phi\left(\gamma \log w_i + Z_i \alpha' - \bar{\psi} \log T_i\right), \tag{1}$$

yielding the standard structural probit specification.[5]

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 in T for any nonzero hours worked. One could try to redefine the reservation wage similarly to before, as 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.[6] The reservation wage is thus set by the following comparison:

- Do not work: then $c = T, 1 - l = 1$, welfare is $\frac{T^{1-\psi}-1}{1-\psi}$.
- Work l^*: then $c = T - \Delta T + wl^*$, $1 - l = 1 - l^*$, welfare is $\frac{(T-\Delta T+wl^*)^{1-\psi}-1}{1-\psi} + \chi\frac{(1-l^*)^{1-\phi}-1}{1-\phi}$.

[5] One could repeat the same exercise using a growth-consistent utility function of the form $\frac{(c \cdot exp(f(1-l)))^{1-\psi}-1}{1-\psi}$. Assuming that $f(1-l) = \frac{(1-l)^{1-\phi}-1}{1-\phi}$, we would get an almost identical probit equation, with an extra constraint of $\gamma = \bar{\psi}$.

[6] 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 the base salary.

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

$$\begin{aligned} \frac{(T+W)^{1-\psi}-1}{1-\psi} + \chi\frac{(1-l^*)^{1-\phi}-1}{1-\phi} &\geq \frac{T^{1-\psi}-1}{1-\psi} \\ \frac{(T+W)^{1-\psi}-1}{1-\psi} - \frac{T^{1-\psi}-1}{1-\psi} &\geq -\chi\frac{(1-l^*)^{1-\phi}-1}{1-\phi}. \end{aligned} \tag{2}$$

One can also give a simple graphical representation (see Figure 3): 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 ($C = T - \Delta T + W, l = l^*$).

To derive a formal expression for the probability of being active (the analogue of (1)), let us linearize the left hand side of (2):

$$\frac{(T+W)^{1-\psi}-1}{1-\psi} - \frac{T^{1-\psi}-1}{1-\psi} \approx WT^{-\psi},$$

so the comparison becomes

$$WT^{-\psi} \geq \chi \underbrace{\frac{1-(1-l^*)^{1-\phi}}{1-\phi}}_{Q} = \chi Q.$$

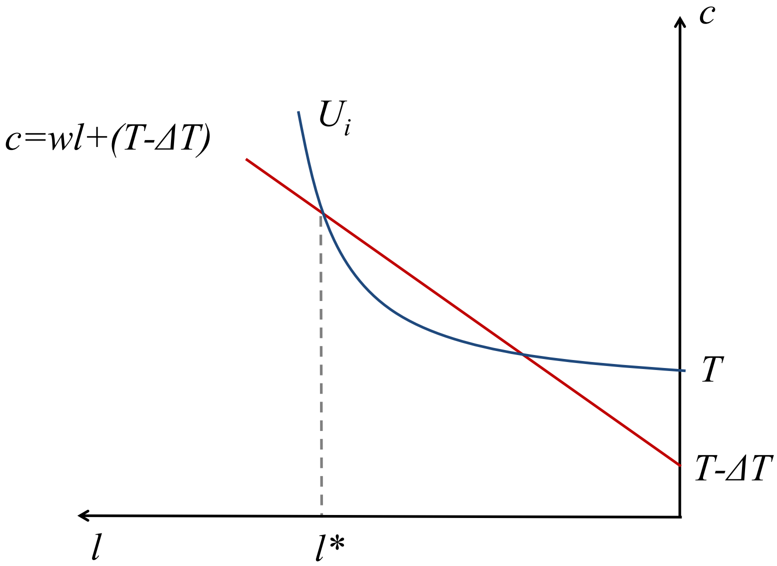
The individual works if

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

yielding again a structural probit of the form

$$P = \Phi\left(\gamma \log W_i + Z_i\alpha' - \bar{\psi}\log T_i\right). \tag{3}$$

Figure 3: The reservation wage when there is a discrete drop in transfers



Let us compare the two structural probit equations (1) and (3). First, W_i in (3) 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 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

Here we closely follow Kimmel and Kniesner (1998), up to a certain point. We want to estimate a structural probit equation:

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

were $W_i = w_i l * -\Delta T_i$. 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 inactives: 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.

In variant A, we thus adopt the following procedure:

1. Run a reduced form probit

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

2. Use the inverse Mills ratio $\lambda\left(x\right) = \frac{\phi(x)}{\Phi(x)}$ 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

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

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

In variant B, we slightly modify the previous procedure:

1. Run a reduced form probit

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

2. Use the inverse Mills ratio $\lambda_i\left(x\right) = \frac{\phi(x)}{\Phi(x)}$ 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^2_W, then one can show that $E\left(log\left(W_i\right)|X_i, Z_i\right) = log\left(E\left(W_i|X_i, Z_i\right)\right) - \frac{1}{2}\sigma^2_W = log\left(e^{X_i\beta_1 + \frac{1}{2}\sigma_1^2} - \Delta T_i\right) -$

κ. 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 (3) 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 other words, 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.[7]

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.[8] 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. Even then, however, it is still the impact of a change in W, not w.

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

$$\frac{\partial \Phi}{\partial \log w} = \frac{\partial \Phi}{\partial \log W}\frac{\partial \log W}{\partial \log w} = \frac{\partial \Phi}{\partial \log W}\frac{w}{w - \Delta T}. \tag{4}$$

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.).

4. Data

We use data from the Hungarian Household Budget Survey (HKF), years 1998-2008. This is in principle a rotating panel database with a one-third renewing part every year,

[7] 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).

[8] 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, making their predicted $\hat{W}_i$ negative.

but it is very difficult to make the actual connections between consecutive waves. For this reason, we only use it as a pooled cross-section. The dataset contains detailed income and consumption measures of broadly $25,000$ individuals per year.

The key challenge is to define the counterfactual transfers: First, how much would someone who is currently working receive in transfers if that individual is laid off? Second, 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 was labour force participation,[9] though we also ran the same estimations with employment. All wage variables (w and W) refer to 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 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 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). In our view, county dummies rep-

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

resent 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 focus mostly on the participation margin: with employment, we only report the results of the main specification but no 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). Table 1 displays our baseline results, following the econometric methodology of Variant A. Panel A reports the estimates for the structural probit equation (3). 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 (4), 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 and look at its employment effect. The resulting 0.28% increase in total employment implies an elasticity of 0.28, quite in line with the consensus.

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.

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.*

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 also much 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 (3)).

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 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 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.*

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.

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.

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 aftertax 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.

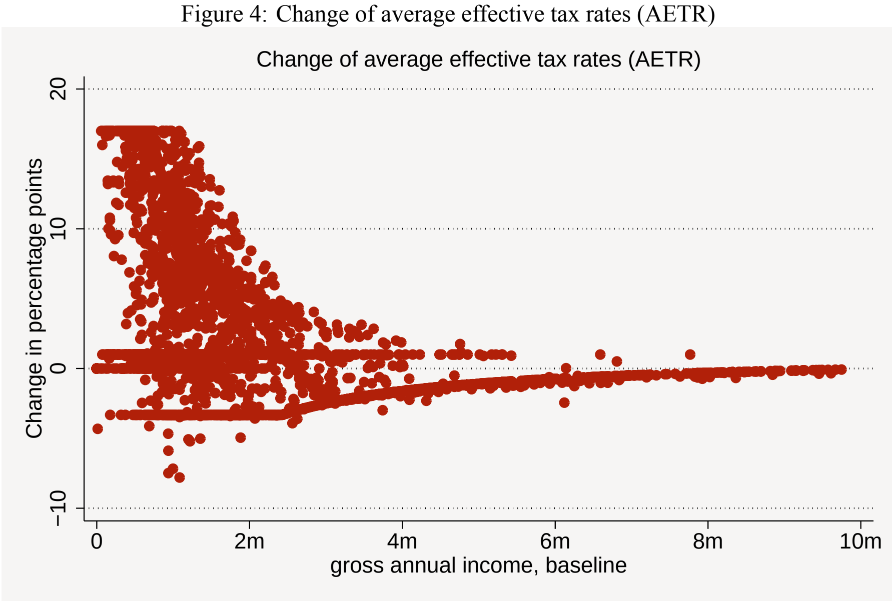
Equipped with these, we form

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

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.

As an illustration, we fed the main changes of the Hungarian personal income tax and transfer system of 2012 into this framework. The particular measures are the following: the complete elimination of the employee tax credit (ETC) scheme, a 27% reduction in the tax rate (from 20,3% to 16%) below the average monthly income of 202,000 HUF, and a 1 percentage point increase in the social contribution rate. As illustrated by Figure 4, 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.

Figure 4: Change of average effective tax rates (AETR)



As a result, aggregate activity decreases by 0.97%, from which the elimination of the ETC is responsible for 2.09%,[10] 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.[11] 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 simultaneously takes into account taxes and transfers. 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.

We find that a single equation can already explain a large heterogeneity of individual responsiveness to taxes and transfers: there are large differences among subgroups, driven partly by a composition effect, and partly by a different share of lost transfers in the GTW.

[10] 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.

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

These highly 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 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)).

References

Aaberge, R., U. Colombino and S. Strom (1999), "Labour Supply in Italy: An Empirical Analysis of Joint Household Decisions, with Taxes and Quantity Constraints", *Journal of Applied Econometrics*, vol. 14, no. 4, pp. 403–22.

Arrufat, J. L. and A. Zabalza (1986), "Female Labor Supply with Taxation, Random Preferences, and Optimization Errors", *Econometrica*, vol. 54, no. 1, pp. 47–63.

Bakos, P., P. Benczúr and D. Benedek (2008), "The Elasticity of Taxable Income: Estimates and Flat Tax Predictions Using the Hungarian Tax Changes in 2005", *MNB Working Papers 2008/7*, Magyar Nemzeti Bank.

Benczúr, P., G. Kátay and Áron Kiss (2012), "Assessing Changes of the Hungarian Tax and Transfer System: A General Equilibrium Microsimulation Approach", *Mnb working papers*, Magyar Nemzeti Bank, forthcoming.

Benczúr, P., G. Kátay, Áron Kiss, B. Reizer and M. Szoboszlai (2011), "Analysis of changes in the tax and transfer system with a behavioural microsimulation model", *MNB Bulletin*, vol. 6, no. 3, pp. 15–27.

Blundell, R. (2012), "Tax Policy Reform: The Role Of Empirical Evidence", *Journal of the European Economic Association*, vol. 10, no. 1, pp. 43–77.

Chetty, R., A. Guren, D. Manoli and A. Weber (2012), "Does Indivisible Labor Explain the Difference Between Micro and Macro Elasticities? A Meta-analysis of Extensive Margin

Elasticities", in: *NBER Macroeconomics Annual 2012, Volume 27*, NBER Chapters, pp. 1–51, National Bureau of Economic Research, Inc.

Dickert, S., S. Houser and J. K. Scholz (1995), "The Earned Income Tax Credit and Transfer Programs: A Study of Labor Market and Program Participation", in: *Tax Policy and the Economy, Volume 9*, NBER Chapters, pp. 1–50, National Bureau of Economic Research, Inc.

Eissa, N. and J. B. Liebman (1996), "Labor Supply Response to the Earned Income Tax Credit", *The Quarterly Journal of Economics*, vol. 111, no. 2, pp. 605–37.

Fernandez, A. I., J. M. Rodriguez-Poo and S. Sperlich (2001), "A note on the parametric three step estimator in structural labor supply models", *Economics Letters*, vol. 74, no. 1, pp. 31–41.

Kimmel, J. and T. J. Kniesner (1998), "New evidence on labor supply: Employment versus hours elasticities by sex and marital status", *Journal of Monetary Economics*, vol. 42, no. 2, pp. 289–301.

Áron Kiss and P. Mosberger (2011), "The elasticity of taxable income of high earners: Evidence from Hungary", *Tech. rep.*

Kátay, G. (ed.) (2009), *Az alacsony aktivitás és foglalkoztatottság okai és következményei Magyarországon [Causes and consequences of low participation and employment in Hungary]*, no. 79 in MNB Occasional Paper, Magyar Nemzeti Bank.

Kátay, G. and B. Nobilis (2009), "Driving Forces Behind Changes in the Aggregate Labour Force Participation in Hungary", *MNB Working Papers 2009/5*, Magyar Nemzeti Bank.

MaCurdy, T. E. (1985), "Interpreting Empirical Models of Labor Supply in an Intertemporal Framework with Uncertainty", in: *Longitudinal Analysis of Labor Market Data*, NBER Chapters, pp. 111–155, Cambridge University Press, heckman, james j. and burton singer ed.

MaCurdy, T. E. (1987), "A Framework for Relating Microeconomic and Macroeconomic Evidence of Intertemporal Substitution", in: *Advances in Econometrics, Fifth World Congress of the Econometric Society*, vol. 2, pp. 149–176, Cambridge University Press, truman f. bewley ed.

Meyer, B. D. and D. T. Rosenbaum (2001), "Welfare, The Earned Income Tax Credit, And The Labor Supply Of Single Mothers", *The Quarterly Journal of Economics*, vol. 116, no. 3, pp. 1063–1114.

Moffitt, R. A. (2002), "Welfare programs and labor supply", in: A. J. Auerbach and M. Feldstein (eds.), *Handbook of Public Economics*, vol. 4 of *Handbook of Public Economics*, chap. 34, pp. 2393–2430, Elsevier.

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),*[12] *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

[12] 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.

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 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.)