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Economic Incentives, Childcare and Gender Identity Norms

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Abstract

This paper investigates the role of gender identity norms in shaping men's and women's time allocation, based on observed behavior following a change in the market penalty for adopting prescriptive norms. To perform this test, we study the reallocation of childcare across parents, following changes in their relative take-home pay. Exploiting variation from Swedish tax reforms, we estimate the elasticity of substitution in parental childcare for natives and immigrants from a variety of countries, characterized by varying gender norms. We find that couples originating from countries with relatively conservative norms are more likely to reallocate childcare across spouses following a reduction in the husband's tax rate, and less likely to reallocate childcare following a reduction in the wife's tax rate, thereby reinforcing a traditional allocation of childcare across parents.

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1 Introduction

The role played by gender identity norms has attracted increasing attention in the study of gender gaps (Bertrand, 2010, 2020). By prescribing appropriate behavior for men and women and inducing utility costs for deviating from underlying norms (Akerlof and Kranton, 2000), gender identity shapes payoffs from several economic actions and potentially feeds into gender gaps in occupational choice, time use and earnings, among other outcomes. As norms typically evolve very slowly, their persistence correlates well with the slowing convergence in gender gaps since the 1990s in the US and other high-income countries, and several studies have detected significant relationships between gender stereotypes and women's aspirations and marital outcomes, with consequences for their labor force participation and earnings.[1]

By the very definition of identity norms, individuals are willing to bear a cost to behave in line with their adopted norms. For example, in a couple that values the male breadwinner model, the wife may underinvest in her career, to the detriment of household income. However, the simple observation of the household and labor market investments of spouses may not be fully informative of their adopted norms, as they typically result from a combination of norms and economic incentives, most notably wage differentials. A natural test of the importance of gender norms would require to observe changes in individual behavior following a change in the market penalty for adopting those norms. For example, a narrowing of the gender wage gap gives couples the incentive to reallocate some of the wife's working time from the household to the market and, conversely, some of the husband's working time from the market to the household. The intensity of such reallocation, which is directly related to the substitutability of spousal inputs in domestic work, is inversely related to the strength of a couple's norms regarding gender roles in the market and the household.

Our paper aims to reproduce this ideal setting in order to investigate behavioral prescriptions regarding the gender allocation of childcare. To this purpose, we combine variation in after-tax wages generated by the introduction of the Earned Income Tax Credit (EITC) in Sweden in 2007, and later EITC adjustments in 2008 and 2009, with administrative information on parental childcare time. The EITC progressively reduced the marginal tax rate

[1] See, among others, Fernandez et al. (2004), Fortin (2005, 2015), Bertrand et al. (2015, 2021), Bursztyn et al. (2017, 2020) and Folke and Rickne (2020).

by about 5 percentage points on low- and mid-level earnings. Thanks to an individual-based tax system, the resulting change in the tax schedule generates independent changes in the marginal tax rates of spouses, according to their respective earnings. Based on these changes, we identify the effect of economic incentives on the spousal division of childcare.

Information on childcare is obtained from administrative sources in the form of Temporary Parental Leave (TPL), available to either spouse to care for a sick child during working hours.[2] Parents are entitled to up to 120 days of TPL per child per year, which is only partly compensated by the social security system, and is recorded in administrative data for social security purposes. According to the Swedish Level of Living Survey, fathers' share of TPL is positively and significantly correlated with fathers' share of overall household work.

Our empirical specification is derived from a household model in which spouses jointly choose their time investments in market work and childcare. Labor supplied to the market earns an after-tax wage determined by the EITC, while spouses' childcare inputs deliver a household public good. We argue that a couple's preferences on how spouses should contribute to the household public good stem from the gender norms it chooses to adhere to.[3] Specifically – and abstracting from technological constraints, to which we come back below – the substitutability of spousal inputs in childcare is the key parameter that captures the strength of norms. Following a change in their respective tax rates, spouses reallocate home (and market) work according to their household optimization problem. For a given tax change, the gain in disposable household income increases with the substitutability between spousal inputs in childcare, which would allow couples to more elastically reallocate their time in line with changed economic incentives. High substitutability indicates that a couple is willing to respond to economic incentives, attaching low importance to specific combinations of spousal inputs' in childcare. Low substitutability indicates instead that a couple has strong preferences regarding inputs' combinations, to the detriment of disposable income.

We estimate that, in our sample of Swedish parents with children aged 4–11, the elasticity

[2]TPL is distinct from Standard Parental Leave (SPL), which is used to care for a baby or young child irrespective of sickness, and is typically used-up before a child's third birthday. The impacts of SPL on household labor supply in Sweden have been studied, among others, by Avdic and Karimi (2018) and Ginja et al. (2020).

[3]Our setting is agnostic about the source of norms, which may reflect intrinsic beliefs and/or social reputation issues. The field experiment by Bursztyn et al. (2020) builds on the difference between these two.

of substitution in childcare is about 1.6. This value is lower than existing estimates of the elasticity of substitution between gender inputs in the labor market, with most values ranging between 2 and 5.[4] However, there is no obvious reason to expect that, in the absence of binding norms, the elasticity of substitution in childcare should be similar to the corresponding estimate for market production, and the use of different methodologies makes it difficult to compare results from the household and the market.

In the absence of a clear benchmark, our discussion of gender norms will hinge on the observed variation in the elasticity of substitution across households in the Swedish resident population. First, we look into variation in the elasticity across groups that potentially differ in their attitudes towards gender roles. Building on the epidemiological approach of Fernandez (2007) and Fernandez and Fogli (2009), which shows evidence that the household formation and work decisions of immigrants in the host country reflect at least in part cultural norms prevailing in their country of origin, we test for heterogeneous responses to tax changes according to norms prevailing in the country of birth of individuals in our sample. As of 2010, about 14.7% of Swedish residents were foreign-born, rising to 19.7% in 2020. The largest immigrant groups are from the Middle East, followed by Eastern and Central Europe, other Nordic countries and Eastern Africa, offering large variation in associated gender norms, as measured by widely used country-level indicators.

Second, we estimate distinct elasticity parameters by exploiting variation from husband's and wife's treatment in turn. The intuition for this distinction is the following. An increase in the husband's after-tax wage, generated by a fall in his tax rate, would induce some labor reallocation towards traditional gender roles, by encouraging him to work more in the market and less in the household, and vice versa for the wife. On the contrary, an increase in the wife's after-tax wage, generated by a fall in her tax rate, would induce labor reallocation away from traditional gender roles. By comparing responses to a fall in husbands' and wives' tax rates across population groups, one can shed light on the importance of "traditional" and "untraditional" gender norms.

We find that non-Nordic immigrant couples tend to react more strongly to a reduction in the husband's tax rate, while the response by Swedish-born or immigrant couples from

[4] See Hamermesh (1993), Weinberg (2000), Acemoglu et al. (2004) and Johnson and Keane (2013).

other Nordic countries is virtually identical for cases of husband's or wife's treatment. We additionally exploit variation across all origin countries represented in our sample by allowing the estimated elasticities to vary continuously with widely-used proxies for gender norms. These include information from gender-related questions in international survey of attitudes, as well as country-level indicators of gender discrimination and gender gaps in relevant outcomes. We find that couples originating from countries with relatively conservative norms are more likely to reallocate childcare across spouses following a reduction in the husband's tax rate, and less likely to reallocate childcare following a reduction in the wife's tax rate. As a result, couples with a more conservative background are more likely to exacerbate gender disparities in childcare time when incentives push in that direction, while they are not as responsive to incentives that would induce a more equal gender division of labor.

The interpretation of our findings hinges on the assumption that the elasticity of substitution between childcare inputs reflects spouses' preferences or beliefs about appropriate gender roles in the household. To rule out alternative channels based on technological substitutability of childcare inputs, or organizational constraints in spouses' respective workplaces, we follow a number of steps. First, we note that technological substitutability is unlikely to drive systematic differences in elasticities that we observe across population groups. Second, we select couples whose youngest child is 4 or older, as mothers may have a biological comparative advantage in the care of younger children. Third, we control in our regressions for parents' specialization in childcare at birth, which may induce comparative advantages in the longer term, and for the workplace composition of mothers and fathers, which would be predictive of a family-friendly work environment and could affect parental contributions to childcare, over and above the impact of gender norms. Finally, in a subsample in which information on occupation is available, we find that our results are robust to the inclusion of mother's and father's occupation and their interaction.

By combining variation from tax reforms and the time allocation of spouses to detect evidence of binding gender norms, this paper contributes to two strands of literature. First, it is related to a recent literature on the role of gender identity norms in the marriage market. Bertrand et al. (2015) have estimated the marriage penalty of deviating from the male breadwinner/female housekeeper model, while Bursztyn et al. (2017) find that single women may avoid career-enhancing actions whenever these signal traits possibly penalized in the marriage market. In our paper, we relate the concept of gender norms to the substitutability of gender inputs in childcare, and design an empirical strategy that allows us to identify this parameter on administrative data. Also, we emphasize variation in identity norms across couples, which we relate to country of origin, rather than across genders. With assortative mating, an important dimension of gender norms may be the variation in decisions and behavior across couples – rather than between men and women – with direct consequences on children's outcomes.

Second, this paper contributes to a broad body of work on the relationship between taxation, labor supply and home production. In the macroeconomic literature, Rogerson (2007) and Ngai and Pissarides (2007) (among others) highlight this relationship as a driver of structural transformation and employment growth, and Bick and Fuchs-Schündeln (2018) relate the labor supply of married couples to variation in tax regimes. The micro literature on labor supply has estimated the impact of taxes on earnings, by exploiting variation from tax reforms and information on taxable earnings from administrative sources (see for example Gruber and Saez, 2002). Closely related to our work, Gelber (2014) estimates the response of spouses' earned income to tax changes in Sweden. Our methodology is similar to Gruber and Saez (2002) and Gelber (2014), and complements their work with a focus on childcare, which offers a direct perspective on gender norms. To our knowledge, this analysis provides the first causal estimates of the impact of taxation on the household division of home production, an effect theoretically studied in Alesina et al. (2011).[5]

The rest of the paper is organized as follows. Section 2 lays out a model of parental childcare and links the elasticity of substitution between spousal inputs to gender norms. Section 3 describes the Swedish institutional background and the data sources. Section 4 builds our empirical framework and discusses identification. Section 5 provides baseline results, Section 6 shows evidence of heterogeneous effects and robustness, and Section 7 concludes.

[5] This paper also adds to a small body of work on the household division of TPL in Sweden. Among others Boye (2015) finds that fathers' TPL take-up is negatively correlated to their contribution to household income, and Ekberg et al. (2013) estimate that it does not respond to their SPL take-up.

2 A model of home production and gender norms

2.1 The couple's optimization problem

Households enjoy a home-produced good H and a market-produced good C and allocate spouses' time between market work and home production. The household good is represented by childcare and produced with a combination of spousal inputs according to the following CES specification:

$$H = \left[sH_m^{\frac{\beta-1}{\beta}} + (1-s)H_f^{\frac{\beta-1}{\beta}} \right]^{\frac{\beta}{\beta-1}}, \tag{1}$$

where $j = m, f$ denotes spouses' gender and $0 \leq H_j \leq 1$ is the share of time devoted by spouse j to childcare. s and $1-s$ are parameters related to the relative efficiency of spouses in childcare and β represents the elasticity of substitution between spousal inputs. We interpret β as representing couples' preferences about the combination of spousal time in childcare and we do not explicitly model technological substitutability between spousal inputs for two reasons. Empirically we can only identify one substitution parameter, hence our interpretation based on preferences implicitly assumes that technological substitutability would not be binding (for example because, from a technological point of view, spousal inputs would be perfect substitutes in the care for 4-11 year olds, in which no biological comparative advantages are involved). Furthermore, we attempt to capture acquired (over and above biological) comparative advantages by controlling for parental specialization after birth, as measured by the shares of standard parental leave taken by each spouse. Finally, evidence on heterogeneous effects presented in Section 6 speaks against a technology-based interpretation of our estimated elasticity of substitution.

In the labor market, the productivity of one unit of time of spouse j is P_j. With perfectly competitive labor markets, wages for each spouse are equal to P_j and the associate earnings Y_j decrease with the share of time devoted to childcare:

$$Y_j = P_j(1 - H_j). \tag{2}$$

Each spouse faces a tax schedule $T(Y_j)$, which may be progressive.

Couples choose the optimal time allocation of spouses in order to maximize joint utility.

Their optimization problem is given by:

$$\max_{H_m, H_f, C} U(H, C) \tag{3}$$

$$\text{s. to.} \quad C \leq [Y_m - T(Y_m)] + [Y_f - T(Y_f)] \tag{4}$$

where H is given by (1) and Y_j is given by (2), with first-order conditions:

$$\frac{\partial U}{\partial H} s H_m^{-\frac{1}{\beta}} H^{\frac{1}{\beta}} = \lambda P_m (1 - \tau_m) \tag{5}$$

$$\frac{\partial U}{\partial H}(1 - s) H_f^{-\frac{1}{\beta}} H^{\frac{1}{\beta}} = \lambda P_f (1 - \tau_f) \tag{6}$$

$$\frac{\partial U}{\partial C} = \lambda, \tag{7}$$

where $\tau_j = T'(Y_j)$ denotes the marginal tax rate and λ is the Lagrangean multiplier. Expressions (5) and (6) represent the compensated demands for childcare inputs, as a function of utility H. At given H, each spouse's contribution to childcare decreases with own labor productivity in the market and net-of-tax income share (NTS), $1 - \tau_j$. The elasticity of demand for childcare inputs with respect to their respective NTS represents the substitution effect of a change in the tax rate, and is given by $\partial \ln H_j / \partial \ln(1 - \tau_j) = -\beta$. Combining (5) and (6) implies that the marginal rate of substitution between spousal inputs in childcare equals the ratio of their opportunity costs:

$$\frac{s}{1 - s} \left(\frac{H_m}{H_f} \right)^{-\frac{1}{\beta}} = \frac{(1 - \tau_m) P_m}{(1 - \tau_f) P_f}. \tag{8}$$

Taking logs on both sides of (8), we obtain

$$h_m - h_f = \alpha + \beta(\sigma_f - \sigma_m) + \beta(p_f - p_m), \tag{9}$$

where lower case letters denote logs, $\alpha = \beta \ln\left(\frac{s}{1-s}\right)$ and $\sigma_j = \ln(1 - \tau_j)$. This expression shows that the relationship between the spousal gap in childcare and the NTS gap hinges on the elasticity of substitution β. As the female-male gap in (log) NTS, $\sigma_f - \sigma_m$, increases with the male-female tax gap, $\tau_m - \tau_f$:

$$\text{sgn}(\beta) = \text{sgn}\left[\frac{\partial(h_m - h_f)}{\partial(\sigma_f - \sigma_m)} \right] = \text{sgn}\left[\frac{\partial(h_m - h_f)}{\partial(\tau_m - \tau_f)} \right].$$

The optimal time allocation can be represented graphically at the tangency between an indifference curve, whose shape is given by equation (1) with $H = \bar{H}$, and a budget constraint representing the opportunity cost of producing $\bar{H}$, whose slope is given by the gender ratio in post-tax wages:[6]

$$K = (1 - \tau_m)P_m H_m + (1 - \tau_f)P_f H_f. \tag{10}$$

This equilibrium is represented by point E^0 in Figure 1, where the wife supplies H_f^0 to childcare, the husband supplies H_m^0, and the cost of home production K can be read on the intercept of the budget constraint on the vertical axis.

2.2 Gender norms and the elasticity of substitution

Let's compare two couples, characterized by different values of β and an equal value of the parameter s. The first couple has $\beta > 0$ and its preferences can be described by the smooth indifference curve in Figure 1. The second couple has instead Leontieff preferences, with $\beta \to 0$, represented by the right-angle indifference curve. In this example we assume that the two couples have an identical time allocation at baseline, represented by point E^0.

Consider a tax reform that reduces the female marginal tax rate, τ_f, and thus increases the gender tax gap, $\tau_m - \tau_f$. The budget constraint becomes steeper and, to achieve the initial level of utility, the time allocation for the first household moves to point E^1, with a reduction in H_f and an increase in H_m. The cost of enjoying a given level of utility has now increased, as illustrated by the higher intercept of the budget line, because the opportunity cost of one of the childcare inputs has increased. The optimal time allocation for the second couple is instead unaffected by the tax change, as spouses are unwilling to alter the combination of spousal inputs due to $\beta \to 0$. This couple faces a higher opportunity cost of childcare than the first household, who is willing to make some input substitution in response to a relative wage change. The higher cost is represented by the ΔK term on the vertical axis.

In our set-up, the differential reaction of the two households to tax changes reflects their different gender identity norms. We argue that the second couple, by not responding to economic incentives, adheres to stricter norms regarding the combination of spousal inputs,

[6] To ease graphical exposition we represent a case of proportional taxation, leading to a linear budget constraint.

and is willing to bear a higher opportunity cost of childcare, according to its preferred combination of spousal inputs. In doing this, the second couple leaves on the table some disposable income, which is instead pocketed by the first couple thanks to its willingness to adjust time allocation in response to economic incentives.

To give a rough sense of magnitudes, one can attempt to evaluate the loss of disposable income, ΔK, following a tax change, for alternative values of β. If one calibrates K to the opportunity cost of TPL, the magnitudes involved are tiny, because couples take on average 7.3 days of TPL per year, which correspond to $7.3/(2\times 251) = 1.5\%$ of their combined working days.[7] But conclusions differ if this framework is applied to the allocation of overall home production time, which amounts to 4 and 4.9 daily hours for fathers and mothers, respectively (SCB, 2012). Using full-time equivalent earnings from the Swedish Wage Structure Statistics, the Online Appendix A provides a back-of-envelope calculation of ΔK for a hypothetical couple with Leontieff preferences. Following a cut in the wife's marginal tax rate in line with the EITC, this couple would bear a 24% higher opportunity cost of home production than a couple with $\beta = 1$, and a 34% higher opportunity cost than a couple with $\beta = 2$, where such opportunity cost evaluated at 2006 wages was about 1,223 Swedish Kronas (SEK) per day (corresponding to approx 150 USD).

In the example of Figure 1, couples have symmetric reactions (or lack thereof) to either an increase or a decrease in the tax gap, following from the assumed production function (1). But it may be realistic to allow for asymmetric adjustment in the time allocation of spouses, depending on whether the husband or the wife experiences a fall in their tax rate. Asymmetric reactions may themselves stem from gender norms. To see this, note that a cut in the wife's tax rate would induce her to work more in the market and less in the household, going against the traditional allocation of labor, while a cut in the husband's tax rate would induce opposite changes and thus reinforce the traditional allocation of labor. The relative strength of these two tax responses may therefore reflect gender norms.

An extreme example is a case in which a couple responds with an elasticity of substitution $\beta > 0$ to a cut in (say) the husband's tax rate, and with a zero elasticity to a cut in the wife's tax rate. This example can be easily introduced in our framework by setting a limit

[7] Year 2006 had 251 working days in Sweden according to https://antal.arbetsdag.se/

to the husband's contribution to childcare, i.e.

$$H = \left[(1-s)H_f^{\frac{\beta-1}{\beta}} + s[\min(H_m, H_m^0)]^{\frac{\beta-1}{\beta}}\right]^{\frac{\beta}{\beta-1}}, \tag{11}$$

where H_m^0 represents the husband's childcare time at baseline. Equation (11) implies that any increase in H_m above H_m^0 would be wasted, hence the couple is not willing to substitute male to female childcare whenever τ_f falls. In this case the couple's indifference curve would coincide with the smooth indifference curve in Figure 1 to the south of E^0 (for $H_m \leq H_m^0$), but would be vertical to the north of E^0 (for $H_m > H_m^0$). Thus the couple would still react to cuts in τ_m, but would not react to cuts in τ_f. Viceversa, a couple that would not increase H_f above a baseline of H_f^0 would react to cuts in τ_f, but not to cuts in τ_m. Whenever we allow for asymmetric adjustment to cuts in τ_m and τ_f we define

$$\beta^+ = \left.\frac{\partial(h_m - h_f)}{\partial(\sigma_f - \sigma_m)}\right|_{d\sigma_f > 0} \quad \text{and} \quad \beta^- = \left.\frac{\partial(h_m - h_f)}{\partial(\sigma_f - \sigma_m)}\right|_{d\sigma_m < 0}.$$

A couple conforms to a traditional gender identity norm if $\beta^+ < \beta^-$; while it conforms to an untraditional gender identity norm if $\beta^+ > \beta^-$.

A potential confounding factor in the interpretation of asymmetric adjustment to cuts in τ_m and τ_f is related to differences in the constraints that spouses face in their respective workplaces: depending on professional roles, one spouse may face higher career costs than the other following work absences. Differences in constraints could induce a kink in the budget constraint in correspondence of the initial allocation E^0 in Figure 1. For example, if the husband experiences a wage cut whenever he spends more than H_m^0 time in the household, the budget line would be flatter to the north of H_m^0, and again the couple would still react to a cut in τ_m, but not to cuts to τ_f. While similar to the previous case, the asymmetric adjustment of the couple's time allocation in this case would be driven by workplace constraints rather than gendered norms. In our empirical analysis we will cater for differences in professional constraints by controlling for workplace characteristics and the occupation of individuals.

2.3 From the model to the data

Equation (9) summarizes the key model result that we bring to the data, allowing us to identify β from the response of the male-female gap in (log) childcare time to changes in the female-male gap in (log) NTS. The steps leading to (9) are based on compensated demands for home inputs (5) and (6), whose elasticity of substitution β measures (the negative of) the substitution effect of a change in the tax rate. Changes in actual (uncompensated) demand, which we measure in the data, also feature income effects of tax changes. The change in the gap in uncompensated home production demands is then given by

$$\Delta(h_m - h_f) = \beta\Delta(\sigma_f - \sigma_m) + \gamma_m\Delta\theta_m - \gamma_f\Delta\theta_f + \beta\Delta(p_f - p_m), \tag{12}$$

where $\Delta\theta_j$ approximates the income effect of a tax change for spouse j and γ_j is the associated parameter. Following Gelber (2014), we measure income effects using each spouse's virtual income, which represents the intersection of the individual's extended budget segment with the y-axis, in a space that has earnings on the horizontal axis and disposable income on the vertical axis. This is the appropriate measure of income effects in the presence of progressive taxation, leading to a nonlinear budget set (Burtless and Hausman, 1978). Changes in virtual income encompass changes in marginal tax rates, changes in (net) benefits and changes in (net) capital income – all of which are taken into account in the construction of $\Delta\theta_j$.

The change in the equilibrium time allocation (12) leads to the following empirical specification for the estimation of the elasticity of substitution β:

$$\Delta(h_{imt} - h_{ift}) = \beta\Delta(\sigma_{ift} - \sigma_{imt}) + \gamma_m\Delta\theta_{imt} - \gamma_f\Delta\theta_{ift} + \gamma X_{it} + (u_{imt} - u_{ift}), \tag{13}$$

in which i denotes couples, t denotes years, X_{it} are observable determinants of the change in the gender gap in market productivity $\Delta(p_{ift} - p_{imt})$, and $(u_{imt} - u_{ift})$ captures unobservable components. To avoid dropping observations for which $H_{imt} = 0$ or $H_{ift} = 0$ we compute the dependent variable in (13) as $\Delta(\ln(H_{imt} + 1) - \ln(H_{ift} + 1))$.[8]

We will discuss the empirical challenges that arise from estimating (13) after describing the data and discussing details of the EITC reform, which will shape our empirical strategy.

[8]Our results are generally robust to other choices, such as adding 0.1 or 0.5 before taking logs.

3 The Swedish institutional setting and data

Sweden represents a valuable testing ground for our approach to the study of gender identity for several reasons. First, the EITC provides exogenous variation in the NTS of spouses and thus in the cost of following gendered norms. Second, Swedish registry data contain longitudinal information on how parents share the care for sick children during their regular working hours, under the TPL program. Although this is only one component of the set of childcare tasks that spouses typically share, it is a good proxy for the gender division of household work, it is measured precisely for the universe of couples on data normally unavailable for other countries, and can be linked to earnings and taxes. Third, while Sweden has one of the highest female employment rates and among OECD countries (Olivetti and Petrongolo, 2017), previous work has found evidence of glass ceiling effects (Breen and Garcia-Penalosa, 2002; Albrecht et al., 2003, 2015), large motherhood penalties (Angelov et al., 2016; Kleven et al., 2019), and higher divorce rates for women who enter politics (Folke and Rickne, 2020). These pieces of evidence may suggest important heterogeneities in the type and strength of gender norms across different population groups. Below we describe the data sources and working sample, the EITC reform and the TPL scheme.

3.1 Data and sample

We use data drawn from several administrative registers compiled by Statistics Sweden, spanning the years 2003–2009.[9] Our primary dataset is the LOUISE register, which covers the resident population of Sweden aged 16–75 and contains information on demographics (age, gender, household composition and country of birth), educational qualifications and a large set of income-related variables. Specifically, it provides information on gross annual labor earnings, government transfers (TPL benefits, Standard Parental Leave (SPL) benefits, sickness insurance, unemployment insurance, etc.) and capital income (interest payments, capital gains on stocks, property sales, etc.).

We match records from the LOUISE register to the multi-generational register, linking

[9]The Social Insurance Database held at IFAU ends in 2010, but we cannot use 2010 data because information on TPL is incomplete.

parents and children, and to the TPL register, compiled by the Swedish Social Insurance Agency, containing start and end dates of TPL spells. We organize our data as a collection of repeated cohorts. In each year $t-1$ we select couples in which both parents are eligible for TPL and participate to the labor force. This defines "cohort $t-1$". We then observe changes in tax treatment and TPL for each spouse between $t-1$ and t. As a consequence, a couple may feature in multiple cohorts as long as it satisfies our selection criteria in the corresponding baseline years. In particular, we select individuals who:

- Live in a couple (married or cohabiting) and have labor earnings above the price base amount (*prisbasbelopp*) in year $t-1$.[10] This restriction ensures that both spouses have sufficient labor market attachment to be potentially affected by TPL entitlement. Non-working individuals are not eligible, while individual working very few hours may not need to rely on TPL to care for a sick child.
- Have their youngest child turning 4–10 in year $t-1$. Parents of younger children may still use SPL for childcare, plus the substitutability between parental inputs in the care of younger children may conflate both biological gender differences and preferences. TPL eligibility expires once a child turns 12, thus the 10 years cap in year $t-1$ ensures that a child's parents are still eligible for TPL when its change is recorded in year t.

Our working sample consists of 463,533 observations who fulfill these criteria. Summary statistics are presented in Table 1. In the representative couple, the husband is 41 years old and the wife is 39 years old, with annual labor earnings of about 386 and 249 thousand SEK respectively. About 11% of individuals are foreign-born. On average, couples jointly take-up 7.3 days of TPL per year, of which 2.4 are taken by the husband and 4.9 are taken by the wife. In 22% of couples, neither spouse takes TPL; in 4% of couples spouses take identical (positive) levels of TPL; in 54% of couples the wife takes more TPL than the husband; and in the remaining 20% of couples the husband takes more TPL than the wife.

[10] For simplicity, we use the terms "spouse", "husband" or "wife" for all cohabiting individuals, irrespective of marital status. The price base amount, is set (and annually revised) by the Government to benchmark welfare benefits. The level was 39,700 SEK in 2006, 40,300 SEK in 2007, and 41,000 SEK in 2008. As of March 2021, the USD to SEK exchange rate is 8.6 and the EUR to SEK exchange rate is 10.2.

3.2 The Earned Income Tax Credit reforms

The EITC was introduced in the Swedish tax system in January 2007 by the centre-right Government coalition, Alliansen, to incentivize labor participation of low- and middle-income earners. The EITC was first highlighted in the Alliansen's election manifesto in August 2006 (Riksrevisionen, 2009), and its design was outlined in October 2006, once Alliansen won the election. The EITC was later reinforced in 2008, 2009, 2010, 2014 and 2019. As our sample ends in 2009, we will consider changes to the tax schedule during 2007-2009.

There are two layers of taxation in Sweden, at the municipal and central government levels. Both municipal and central government taxes are levied on individual taxable earnings, given by gross earnings minus a basic deduction. The EITC introduced additional deductions for low- and middle-income earners that were solely a function of earnings, and unrelated to marital status or the presence of qualifying children in the household. An individual's tax liability is automatically reduced by the tax credit, if eligible, thus take-up is universal. After being emphasized as one of the Alliansen's flagship policies, the incidence of the EITC was salient to employees from their first 2007 pay slip.

Municipal tax rates are proportional and equal on average to 31.6% in 2006 (ranging from 28.89%–34.24% across municipalities). The central government tax rate is progressive and is phased in at 20% for earnings above 306,000 SEK, rising to 25% for earnings above 460,600 SEK. The solid line of Figure 2 plots marginal tax rates in 2006 for a representative municipality with a local tax rate of 31.6%. Earnings below 306,000 SEK are only subject to municipal taxes but, due to a complex system of deductions, the tax schedule between 0 and 306,000 SEK has the irregular shape represented by the solid line in Figure 2. Beyond 306,000 SEK, the combination of municipal and central government taxes yields a 51.6% marginal tax rate up to 460,600 SEK, and 56.6% thereafter.

The 2007 EITC introduced additional deductions that lowered the marginal tax rate in two intervals of the earnings distribution, as represented by the blue dashed line in Figure 2.[11] For very low earnings between 17,000–32,000 SEK, the marginal tax rate was cut

[11] The Online Appendix B gives further detail on the incidence of the EITC along the distribution of taxable earnings and shows how marginal tax rates are computed before and after its introduction. See also Edmark et al. (2016) and DalBo et al. (2020) for descriptions of the EITC.

to 0. For intermediate earnings between 123,500–306,000 SEK, the marginal tax rate fell by 3.2 percentage points on average (2.9–3.4 points across municipalities), from 34.8% to 31.6%. Based on our selection criteria, we do not exploit variation from large tax changes for very low earners, but only from small tax changes for middle earners. Given the earnings distribution for men and women at the time, the 2007 treatment range [123,500–306,000 SEK] is centered around median female earnings of 218,100 SEK and lies below median male earnings of 320,000 SEK. In 2008, additional deductions reduced the marginal tax rate to 30.6% in the representative municipality for earnings in the range 109,600–282,100 SEK, as shown by the dashed green line in Figure 2 and in 2009 marginal tax rates further fell to 29.5% for earnings in the same range (dashed red line). Below this range, the 2009 EITC further reduced the marginal tax to 23.7% for earnings between 37,300–109,600 SEK.

To have a sense of changes in the unit cost of TPL, consider the following example, based on the 2007 EITC. A person with earnings in the intermediate range 123,500–306,000 SEK has her marginal tax rates reduced from 34.8% in 2006 to 31.6% in 2007. Before the EITC, she would give up 13% of her daily income for taking one day of TPL. After the EITC, she would give up 18.3% of her daily income.[12] This 25% increase in the opportunity cost of TPL may induce the household to reallocate TPL from the treated spouse to the untreated one. On the other hand, given that the average couple takes 7.3 days of TPL a year, the change in the opportunity cost of TPL represents a negligible component of household earnings, leaving several dimensions of household finances largely unaffected (wealth, consumption smoothing, spousal bargaining power, etc.), and we would not need to model these dimensions explicitly in our analysis.

Given individual taxation, couples may face a higher, lower or unchanged tax gap, depending on spouses' baseline earnings. To measure treatment, we compute changes in simulated tax rates, given by the tax change an individual would experience at constant earnings, unaffected by endogenous labor supply responses to the EITC. These are computed by ap-

[12] A person with daily earnings equal to y would give up $(1-0.348)y-0.8(1-0.348)y = 0.130y$ without the EITC, where 0.8 represents the (uncapped) TPL replacement rate, and $(1-0.348+0.032)y-0.8(1-0.348)y = 0.162y$ with the EITC.

plying the tax schedule in year t to $t-1$ earnings,

$$\Delta\tilde{\tau}_{ijt} = \tilde{\tau}_{ijt} - \tau_{ijt-1} = T'_{ijt}(Y_{ijt-1}) - T'_{ijt-1}(Y_{ijt-1}) \tag{14}$$

and can be used to construct simulated changes in NTS, $\Delta\tilde{\sigma}_{ijt} = \ln(1-\tilde{\tau}_{ijt}) - \ln(1-\tau_{ijt})$.

Figure 3 shows the distribution of changes in simulated tax gaps. Pooling together 2007–2009 tax changes, 46.5% of observations, represented by the central spike in the distribution, experience no change in their tax gap $\tau_{imt} - \tau_{ift}$ and thus constitute our control group. Figure 4 shows examples of control couples in the 2007 reform: these include cases in which neither spouse is treated (e.g. the wife earns Y_1 and the husband earns Y_4 or viceversa) and cases in which spouses are affected in the same way (e.g. the wife earns Y_2 and the husband earns Y_3 or viceversa). The remaining 53.5% of couples are treated. Among these, 85.5% experience an increase in the tax gap of up to 3.4 percentage points (5.3% change in log NTS), following a reduction in the female tax rate. Wives in these couples have earnings in the treated region (say Y_2 or Y_3 in Figure 4), and husbands have earnings outside this region (Y_1 or Y_4). Finally, 14.5% of couples experience a symmetric reduction in the tax gap, of up to 3.4 percentage points (5.3% change in log NTS), following a reduction in the male tax rate. Husbands in these couples have earnings in the treated region, and wives have earnings below or above.

Variation in tax treatment across the 2007–2009 reforms will be exploited in Section 4 to identify the elasticity of substitution between spousal inputs in childcare.

3.3 Temporary Parental Leave

TPL can be used by either parent to care for a sick child aged between 8 months and 12 years. While representing only a portion of overall home production, TPL take-up is directly linked to parenthood and work-life balance considerations, which are one key driver of gender gaps in earnings (Kleven et al., 2018, 2019). By linking administrative information on TPL spells to survey based information on time use from the 2000 Level of Living Survey (LNU), Eriksson and Nermo (2010) show that fathers' share of TPL is positively correlated to their household work. We replicate this result on a sample of 2000 and 2010 LNU respondents, selected using the same criteria of our main sample, i.e. dual earner couples with children aged 4 or above, and find that fathers' TPL days are positively and significantly correlated to their share of household work, having controlled for total spouses' home production time and their human capital levels.[13] Similarly as Ekberg et al. (2013), we use the parental division of TPL as a proxy for the parental division of childcare.

TPL is compensated at 80% of foregone earnings up to a cap of 302,250 SEK in 2007, 307,500 SEK in 2008 and 321,000 SEK in 2009. Parents are jointly eligible for a maximum of 120 days per child per year, though they are not allowed to take TPL at the same time, with minor exceptions in cases of major illnesses and/or hospitalization.[14] To receive TPL benefits, a parent needs to register a child as sick on the first day of the sickness spell, and from the eighth day a doctor's certificate is required. While the contemporaneous cost of TPL to employees is arguably small, it represents the only component of home production that is clearly visible to employers, and may signal patterns of spousal specialization in home production and shape employers' beliefs about productivity and job attachment.

The use of TPL in Sweden is widespread. In 2016, 458,260 mothers and 353,289 fathers used TPL, respectively, for a total of over 6.3 million days. In our sample, men and women take on average 2.4 and 4.9 TPL days per year, respectively, and 22% of couples have positive TPL (see Table 1). While TPL represents only a small reduction to individuals' labor supply, its take-up is at an employee's discretion and is less likely to be constrained by optimization frictions that typically interfere with working hours adjustment (see, among others, Chetty, 2012). Importantly, there seems to be little or no competition to TPL use from the market childcare sector in Sweden. As of 2016, only 1.4% of private expenditure for household services was accounted for by childcare services (with cleaning and gardening jointly accounting for 94.7%).[15] This amounts to about 103 million SEK of private expenditure on childcare, against 146 billion SEK of Government expenditure.

[13] Our estimated coefficients imply that a ten-percentage point rise in fathers' share of total home production is associated with one additional half-day of father's TPL.

[14] TPL can be taken for a full day or 75%, 50%, 25% or 12.5% of a day. If parents are unable to provide care, for instance because they are sick themselves, TPL eligibility can be transferred to a third person.

[15] Figures based on tax deductions of individuals purchasing household services on the market.

4 The econometric framework

We bring specification (13) to the data described above. Note that the error term in (13), $u_{imt} - u_{ift}$, is likely to be correlated with the the substitution effect $\Delta(\sigma_{ift} - \sigma_{imt})$ and income effects $\Delta\theta_{ijt}$, as it absorbs unobserved components of the change in the productivity gap $\Delta(p_{ift} - p_{imt})$. Under progressive taxation, $\Delta(p_{ift} - p_{imt})$ is negatively correlated to $\Delta(\sigma_{ift} - \sigma_{imt})$, as higher productivity feeds into higher incomes and marginal tax rates. For the same reasons, $\Delta(p_{ift} - p_{imt})$ is negatively correlated to $\Delta\theta_{ift}$ and positively correlated to $\Delta\theta_{imt}$. In general, the marginal tax rate faced by an individual is endogenous to their labor supply choices, which may in turn be affected by the tax reform.

The EITC generated exogenous changes in spouses' marginal tax rates that can in principle be used to obtain valid instruments for changes in NTS and income effects. Conditional on base-year earnings and pre-treatment characteristics of spouses, the change in the simulated NTS gap, $\Delta(\tilde{\sigma}_{ift} - \tilde{\sigma}_{imt})$, provides a potential instrument for the change in the actual NTS gap $\Delta(\sigma_{ift} - \sigma_{imt})$ in equation (13). Similarly, simulated income effects $\Delta\tilde{\theta}_{ijt} = \tilde{\theta}_{ijt} - \theta_{ijt-1} = \ln[Y_{ijt-1} - T_{ijt}(Y_{ijt-1})] - \ln[Y_{ijt-1} - T_{ijt-1}(Y_{ijt-1})]$ could be used as instruments for actual income effects $\Delta\theta_{ijt}$. This strategy, however, opens up two challenges, discussed below.

4.1 Monotonicity

The first issue is that simulated tax changes from the EITC are not everywhere monotonically related to actual tax changes in our setting. In fact we find that individuals whose earnings in year $t-1$ are just below the earnings threshold for treatment in year t on average experience some earnings growth between $t-1$ and t. Some of these individuals, who would ex-ante be eligible for a *cut* in their marginal tax rate via the EITC, move across the treatment threshold, and thus ex-post experience an *increase* in their actual tax rate. Vice versa for earnings just above the threshold.

This issue is illustrated for the year 2007 in Figure C–1 of the Online Appendix, which plots actual and simulated marginal tax changes by 10,000 SEK bins of the earnings distribution. Panel A refers to men and Panel B refers to women. Grey dots indicate simulated tax changes, and white and black dots indicate actual tax changes during 2005–06 and 2006–07, respectively.[16] For both men and women, in a few bins below the 306,000 SEK threshold, individuals would have experienced a reduction in their tax rate had their income remained at the 2006 level (grey dots), but experienced instead an increase in their actual tax rate due to their increased earnings (black dots). Vice versa, just above the same threshold, the change in the simulated tax rate is null but the actual change is negative. Note that analogous discrepancies between actual and simulated tax changes can be observed in the vicinity of each taxation threshold, and they are not specific to the reform year but are also present in 2005–06 (as shown by the white dots). As one would expect, the vertical distance between the 2006–07 and 2005–06 lines is positive where the simulated tax falls, and tends to zero in the other relevant intervals of the earnings support.

Given the violation of the monotonicity assumption in some ranges of the earnings distribution, we identify the impacts of interest from the following intention-to-treat specification, in which we regress the change in the TPL gap on (functions of) simulated tax changes:

$$\Delta(h_{imt} - h_{ift}) = \beta\Delta(\tilde{\sigma}_{ift} - \tilde{\sigma}_{imt}) + \gamma_m \Delta\tilde{\theta}_{imt} - \gamma_f \Delta\tilde{\theta}_{ift} + \gamma X_{it} + (\tilde{u}_{imt} - \tilde{u}_{ift}). \qquad (15)$$

4.2 Controlling for the counterfactual evolution of the TPL gap

Estimation of (15) identifies the causal effects of interest if, conditional on an appropriate set of controls, the assignment of couples to alternative tax regimes can be considered as good as random, i.e. they would experience a similar evolution of the TPL gap in the absence of a tax reform. While the inclusion of rich controls for base-year earnings may adequately proxy for the counterfactual evolution of the TPL gap, such controls may impair identification in our context, as tax variables on the right-hand side are themselves deterministic functions of earnings (see also Gruber and Saez, 2002 and Gelber, 2014 for a similar discussion). In the absence of earnings' controls, the error term in (15) embodies the counterfactual change in the TPL gap, $\Delta(h_{imt} - h_{ift} | \text{no reform})$, and the resulting bias in the β estimate depends on its correlation with the regressor $\Delta(\tilde{\sigma}_{ift} - \tilde{\sigma}_{imt})$.

[16] 2005–06 changes are obtained on a sample selected on the same criteria as those used for our main 2006–07 sample.

To illustrate possible patterns of correlation, let's consider first couples in which the husband earns more. Given that the EITC predominantly applies to mid-low earnings ranges, wives in these couples are more likely to be treated than husbands, i.e. $\Delta(\tilde{\sigma}_{ift} - \tilde{\sigma}_{imt}) > 0$. These wives take up the bulk of TPL at baseline, $h_{imt-1} - h_{ift-1} < 0$, but (other things equal) such differential tends to shrink over time because (i) total TPL declines with average children's age and (ii) this decline mostly bites on the TPL share of the main provider, implying $\Delta(h_{imt} - h_{ift}|\text{no reform}) > 0$.[17] For a symmetric argument, in female breadwinner couples, the husband is both more likely to be treated and to reduce his TPL contribution when children age, thus $\Delta(\tilde{\sigma}_{ift} - \tilde{\sigma}_{imt}) < 0$ and $\Delta(h_{imt} - h_{ift}|\text{no reform}) < 0$ in this case. The positive correlation between $\Delta(\tilde{\sigma}_{ift} - \tilde{\sigma}_{imt})$ and $\Delta(h_{imt} - h_{ift}|\text{no reform})$ would lead to an upward bias in our estimate for β. These are only likely scenarios, and other scenarios are of course possible. In general, if the counterfactual change in the TPL gap is systematically correlated to the baseline earnings gap, which in turn determines spousal treatment, one cannot recover an unbiased estimate for β without controlling for the counterfactual evolution of the TPL gap.

To factor in differences in the evolution of the TPL gap between treated and control couples that would have occurred absent any tax reform, we follow two alternative approaches. The first approach, which we adapt from Gelber (2014), consists in estimating $\Delta(h_{imt} - h_{ift}|\text{no reform})$ from a parametric relationship between TPL and income for the 2005–06 cohort:

$$\Delta(h_{ij06}) = g(y_{if05})\xi_y^{jf} + g(y_{im05})\xi_y^{jm} + X_{if05}\xi_x^{jf} + X_{im05}\xi_x^{jm} + v_{06},$$

where y_{ij05} denotes 2005 log income (earnings plus benefits), $g(.)$ denotes a ten-piece spline (with knots at the deciles), and X_{ij05} are observable characteristics.[18] The estimated coeffi-

[17] Table C–1 in the Online Appendix shows evidence on these patterns before 2007 (when no tax change occurred). Panel A shows that both combined TPL and the TPL gap decline with the age of the youngest child. Panel B shows that the overall decline in TPL is stronger for women than for men as well as for the secondary earner and for the main TPL provider in a couple.

[18] These are: age fixed effects, education fixed-effects (7 categories), dummy for born in Nordic country, industry fixed effects (10 categories) for each spouse; municipality fixed-effects (289); fixed effects for the number of children aged 4-11; fixed effects for the age of the youngest child; total days of SPL taken by the couple; share of SPL taken by the mother; fixed effects for the number of grandparents living in the same municipality; share of mothers with children aged 0-11 at each spouse's workplace (and their interaction); marital status.

cients $\hat{\xi}_y^{jf}$ and $\hat{\xi}_y^{jm}$ calibrate the evolution of TPL in the absence of tax changes at different parts of the income distribution and are used to predict TPL changes for each spouse in later years:

$$\Delta_G(h_{ijt}|\text{no reform}) \quad = \quad g(y_{ift})\hat{\xi}_y^{jf} + g(y_{imt})\hat{\xi}_y^{jm}, \tag{16}$$

where the G subscript denotes the Gelber (2014) procedure.

The second approach builds on a non-parametric prediction of $\Delta(h_{imt} - h_{ift}|\text{no reform})$, obtained by matching couples in our sample to couples in the pre-reform period. For each spouse j in couple i in the sample, we select the S_{ij} set of spouses in 2005, who falls in the same bin (of size 1/25) of the earnings distribution and live in the same municipality. We then predict the change in TPL between $t-1$ and t in the absence of tax reforms based on the change in matched couples during 2005–06:

$$\Delta_M(h_{ijt}|\text{no reform}) = \mathbb{E}\{\Delta(h_{ij06}|ij \in S_{ij}\}, \tag{17}$$

where the subscript M stands for the matching procedure.

Based on either counterfactual (16) or (17), we construct $\Delta_k(h_{imt} - h_{ift}|\text{no reform})$, $k = G, M$, which we use to residualize the dependent variable in (15) and estimate:

$$\hat{\Delta}_k(h_{imt} - h_{ift}) = \beta\Delta(\tilde{\sigma}_{ift} - \tilde{\sigma}_{imt}) + \gamma_m\Delta\tilde{\theta}_{imt} - \gamma_f\Delta\tilde{\theta}_{ift} + \gamma X_{it} + \epsilon_{it}, \tag{18}$$

where $\hat{\Delta}_k(h_{imt} - h_{ift}) \equiv \Delta(h_{imt} - h_{ift}) - \Delta_k(h_{imt} - h_{ift}|\text{no reform})$.

With either residualization method, this procedure amounts to a triple difference identification strategy, in which differences across control and treated couples before and after the EITC are benchmarked against the corresponding differences in a period without tax changes. Identification of the effects of interests requires that a parallel trends assumption is satisfied. That is, in the absence of tax reforms, the evolution of the change in the TPL gap should differ at most by a constant between treated and control couples.

4.3 Asymmetric responses

We finally allow for asymmetric responses in a couple's TPL gap to an increase or decrease in the tax gap, generated by a reduction in the wife's or husband's tax rates, respectively, in the following variant of equation (18):

$$\begin{aligned}\hat{\Delta}_k(h_{mt} - h_{ft}) &= \beta^+ \Delta(\tilde{\sigma}_{ft} - \tilde{\sigma}_{mt})|_{\Delta\tilde{\sigma}_{ft}>0} + \beta^- \Delta(\tilde{\sigma}_{ft} - \tilde{\sigma}_{mt})|_{\Delta\tilde{\sigma}_{mt}>0} \\ &+ \gamma_m \Delta\tilde{\theta}_{imt} - \gamma_f \Delta\tilde{\theta}_{ift} + \gamma X_{it} + u_t, \end{aligned} \tag{19}$$

by imposing a spline in $\Delta(\tilde{\sigma}_{ft} - \tilde{\sigma}_{mt})$, with a knot at 0.

5 Average elasticity estimates

We first show separate estimates for each reform year, as well as for the pre-reform period, which serves for falsification. The vector of couples' characteristics X_{it} includes individual and household demographics, together with education and industry effects for each spouse. Importantly, we also control for several factors potentially shaping the couple division of childcare, over and above the significance of gender norms.[19] First, as mothers' comparative advantage at birth may lead them to specialize in childcare and become less substitutable in the care for older children, we introduce controls for parents' specialization at birth, as proxied by the share of SPL taken by the mother for all children in the household. This share is included together with the total amount of SPL jointly taken by parents.

Second, another factor that could possibly affect the substitutability between parents' childcare time is the presence of respective workplace constraints. The law leaves discretion to employees on the use of TPL, but individuals may take into account disruption effects generated by their work absences and potential career consequences. While we have no direct information on family-friendly workplace practices, we use as proxies of a workplace culture the share of female employees with young children (and thus eligible for TPL) in each spouse's workplace and the interaction between the two. A complementary analysis consists in adding occupation controls, as occupations may capture patterns of substitutability at

[19] Estimates that do not control for these factors are not reported (available on request), but they are extremely close to those discussed below.

work. Information on occupation is only available for a 30% subsample of employees and we will present results on this reduced sample in the robustness analysis of Section 6.3.

Third, we noted in Section 3.3 that the private sector provides scant alternatives to TPL, as shown by the negligible take-up of tax deductions related to households' private expenditure on childcare services. However, this may not rule out the use of informal child care. Based on the Swedish Multi-generational Register we build a proxy for the most common type of informal childcare, provided by grandparents who live within commuting distance. We thus link spouses in our sample to their parents (if present in the data) and control for the number of grandparents living in the same municipality.

Estimated elasticities for each cohort are shown in Figure 5. Estimates for the 2006-07, 2007-08 and 2008-09 cohorts (our working sample) are based on specification (18). Changes in TPL gaps between $t-1$ and t have been residualized using corresponding changes in the 2005-06 cohort and the Gelber (2014) method (Panel A) or the matching method (Panel B). Estimates for the 2004-05 and 2005-06 cohorts (our falsification period) are obtained after residualizing changes in TPL gaps between $t-1$ and t using corresponding changes in the 2003-04 cohort. For the 2004-05 and 2005-06 cohorts we impose the 2007 hypothetical treatment range.[20] With either residualization method, pre-treatment estimates are fairly precise zeros, while post-treatment estimates are positive and highly significant. Estimates during the reform years range from 1.38 for the 2006-7 cohort to 2.80 for the 2008-09 cohort in Panel A and from 1.40 to 3.36 in Panel B. There is no obvious interpretation for the rise in the estimated elasticities over time. One possible explanation is that the impact of the tax treatment is heterogeneous across the earnings distribution – as tax changes during 2008-09 affect earnings further down in the distribution relative to changes in earlier years. Another explanation is that individuals may become progressively better aware of the EITC, which triggers more widespread responses.

Table 2 shows estimates on the pooled EITC sample in columns 1-3 (2006-07, 2007-08 and 2008-09 cohorts) and for the placebo sample in columns 4-6 (2004-05 and 2005-06 cohorts). To highlight the role of residualization, we show both double-difference estimates (columns

[20] Imposing the 2008 or 2009 treatment ranges for the 2004-05 and 2005-06 cohorts instead of the 2007 treatment range yields equivalent placebo estimates, both qualitatively and quantitatively.

1 and 4, based on specification (15)) and triple-difference estimates (columns 2-3 and 5-6, based on specification (18)). Each column in Panels A and B is obtained from a separate regression. In column 1 of Panel A, the dependent variable is the raw change in the (log) TPL gap, and the resulting β estimate is about 2.2 and highly significant. In column 2 the change in the TPL gap is residualized with respect to its counterfactual evolution based on the Gelber (2014) method, and the β estimate falls to about 1.6, which corresponds to the average of the three separate estimates reported in the left panel of Figure 5. A similar estimate is obtained in column 3, in which the dependent variable is residualized using the matching method. As one may have expected from the discussion of Section 4.2, the double-difference estimate for β in column 1 is upward biased with respect to estimates in columns 2 and 3 that control for the counterfactual evolution of the TPL gap.

Columns 4-6 report results from the corresponding placebo regressions. We detect a significant estimate in column (4) on the double-difference specification, implying that couples whose earnings lie in the treated ranges – but receive no treatment – display systematically different TPL behavior from other couples in the absence of tax reforms,[21] and, only when these differences are catered for with the Gelber (2014) or matching adjustments in columns (5) and (6) respectively, is the placebo estimate very close to zero and statistically insignificant.

Panel B allows for asymmetric responses to changes in tax gaps generated by wives' and husbands' tax cuts, which identify β^+ and β^-, respectively, as in specification (19). In column 1, based on the double-difference specification, the response of the TPL gap to a decrease in the female tax rate is significantly smaller than its response to a decrease in the male tax rate, but this difference is much reduced and not-significantly different from zero once the dependent variable is residualized in columns 2 and 3. We detect again significant placebo effects when the dependent variable is not residualized in column 4, and these vanish with either the Gelber (2014) or matching adjustments in columns 5 and 6, respectively.

In summary, the estimates presented in Table 2 suggest an elasticity of substitution around 1.6, with no significant variation across husbands' and wives' treatment or residualization methods. Importantly, all placebo estimates based on residualized dependent

[21] This was also noted by Edmark et al. (2016) in double-difference estimates of the EITC on employment.

variables are not statistically significant and very close to zero in magnitude.

The literature contains a number of estimates for the elasticity of substitution between male and female inputs in the labor market. Hamermesh (1993) reviews various studies and suggests values of such elasticity of 2 for the U.K. and 2.3 for Australia. For the U.S., Weinberg (2000) and Acemoglu et al. (2004) obtain estimates around 2.4 and 3, respectively, and Johnson and Keane (2013) obtain a higher estimate just above 5.[22] These and our estimates are obtained on different methodologies and contexts, but – with these qualifications in mind – the comparison between them hints that the substitutability of gender inputs in market work is overall higher than their substitutability in childcare.

6 Heterogeneous effects by country of origin

6.1 Indicators of gender norms

We next consider heterogeneity in spousal responses to tax incentives by country of origin, leveraging both variation in behavior between natives and immigrants, and among immigrants from different cultural backgrounds.[23] In doing this we build on the epidemiological approach, which identifies the role of culture from variation in behavior across individuals observed within the same economic and institutional environment, but whose norms are potentially different.[24]

To measure cultural, economic and legal differences in gender-related matters between couples from different countries, we consider a number of indicators often associated to cross-country variation in gender norms. Some of the indicators we use are based on direct elicitation of norms, while others summarize institutional or economic aspects that can be

[22]A common approach in this literature consists in regressing the log gender wage gap in the aggregate economy (or partitions within it) on the gender gap in log labor supply, having imposed a trend increase on the relative demand for female labor. Acemoglu et al. (2004) use individual wages and instrument relative female labor supply in 1940s US with male mobilization rates.

[23]Clearly, individuals originating from different countries are also likely to differ in several other dimensions (most notably earnings and the counterfactual evolution of the TPL gap in the absence of tax changes) that are related to the outcomes of interest. The role of unobservables at the couple level is factored in by taking first differences. In addition, we residualize the dependent variable with respect to the predicted counterfactual TPL gap, as described in Section 4.2 and include a rich set of controls X_{it}.

[24]See Fernandez (2011) for applications of this approach to female labor supply and fertility decisions, as well as applications to other contexts.

associated to variation in gender norms. From the World Value Survey (waves 5-7), we use responses to the statement: "When jobs are scarce, a man should have more right to a job than a woman". This is available for the largest set of countries and is widely used as a measure of gender norms (see e.g. Azmat et al. 2006, Alesina et al. 2013 and Bertrand et al. 2021). We use an additional statement from the World Value Survey, "When a mother works the children suffer," referring more specifically to spouses' contributions to childcare, which are the focus of our analysis. We use the percentage of the population that disagrees with each statement (denoted by WVS1 and WVS2, respectively) as indicator of liberal gender norms. In addition, we use global indices relating to gender equality, namely the Global Gender Gap Index (GGGI), capturing gender gaps in economic participation and opportunities, educational attainment, health, survival, and political empowerment; and the Social Institution and Gender Index (SIGI), which captures discrimination in social institutions, aggregating sub–scores that relate to women's discrimination in the family, restricted physical integrity, restricted access to productive and financial resources, and restricted civil liberties. Finally, we use two economic outcomes that tend to be associated to gender equality, namely the ratio of female to male labor force participation rates (FLFP) and GDP per head. To ease interpretation and discussion, all indicators are defined such that *higher values correspond to higher gender equality* (see also notes to Table 3).

There is wide cross-country variation in all indicators considered. To give an example, the percentage of respondents who disagree with the statement "When jobs are scarce, a man should have more right to a job than a woman," (WVS1) has an unweighted world mean of 50 (corresponding e.g. to Colombia) and a standard deviation of 21, with Sweden at the 95th percentile (83) and Iran at the 5th (12). Pairwise rank correlations between the six indicators range from 0.14 (GDP–FLFP) to 0.75 (SIGI–WVS1).

Table 3 shows descriptive statistics for these indicators in our sample, obtained by assigning to each couple the value of the indicator corresponding to the spouses' country of origin. If spouses were born in different countries, the couple is assigned the mean value of each indicator. As the variable denoting country of birth in the Swedish registry data is aggregated into 27 groups (described in Table C–2), we first compute the indicator for each group as the weighted average of the corresponding country-level indicators, using the shares of migrants in 2000 from each country as weights (available from Statistics Sweden). As one would expect, variation in the indicators is much reduced in the couple-level data, as in 83% of couples both spouses were born in Sweden, but the results presented below imply that our empirical setting has enough statistical power to identify the impacts of interest.

One key divide in the international variation of indicators is between Nordic countries (Sweden, Norway, Finland, Denmark, Iceland) and the rest of the world, as shown in Figure 6. While indicators differ in the way they are constructed, the aspects of gender inequality they reflect, and the purpose for which they are originally obtained, Nordic countries provide a much more favorable ground to gender parity according to any of them.

6.2 Results

We first show separate elasticity estimates for couples in which at least one spouse was born in a Nordic country and couples in which both spouses were born elsewhere. The estimates, reported in Table 4, are obtained in a single regression that includes an interaction between the tax treatments and a dummy for Nordic origin. Columns 1 and 2 in Panel A show similar estimates for the overall elasticity of substitution, and the corresponding difference is not statistically significant (column 3).

Panel B fits different slopes for wives' and husbands' tax cuts and shows that the β^+ and β^- estimates are nearly identical for Nordic couples, who therefore react with the same intensity to reductions in wives' and husbands' marginal tax cuts. There is instead evidence of a wide gap in the elasticities for non-Nordic couples, whose reaction to wives' treatment is weak and not significantly different from zero, while their reaction to husbands' treatment is very large, statistically significant, and more than twice as large as the corresponding elasticity for Nordic couples (3.6 versus 1.6). Horizontal cross-country differences in β^+ and β^- are statistically significant (column 3), while vertical within-group differences between β^+ and β^- are only significant for the non-Nordic couples (column 2). The positive and statistically significant double difference at the bottom of column 3 implies that non-Nordic couples display, overall, a more conservative behavior than Nordic couples, because they respond less intensively to female tax cuts and more intensively to male tax cuts.

We next exploit variation across all origin countries represented in our sample by allowing the estimated elasticities to vary continuously with proxies for gender norms in each couple's country of origin. We standardize each indicator so that estimates can be interpreted as the effect of a one standard deviation change. Standard deviations are reported in Table 3 for reference.

Columns 1-6 in Table 5 report separate estimates for each indicator. Panel A includes an interaction term between the tax treatment received by couples and each indicator in turn, together with its main effects. As indicators are standardized, β estimates represents the elasticity for the average couple and, as expected, the estimates in columns 1-6 coincide almost exactly with the value reported in column 2 of Table 2. The coefficient on the interaction term measures the change in the elasticity associated to a one standard deviation increase in gender equality in the spouses' country of origin. In all specifications this coefficient is small and statistically insignificant, except in column 2 where it is significant at the 10% level. Small estimates on those interaction terms echo the small and statistically insignificant difference between the coefficients in columns 1 and 2 of Panel A in Table 4.

Panel B of Table 5 (similarly as Panel B of Table 4) shows instead that norms in the origin country do matter for asymmetric responses to wives' and husbands' treatment. Specifically, β^+ estimates increase and β^- estimates decreases with a more egalitarian environment, across all indicators considered. The interpretation is that more egalitarian norms boost the response of spouses' time allocation to a cut in the wife's tax rate, while they dampen the response to a cut in the husband's tax rate. The size of the estimated interaction coefficients is remarkably stable across specifications. On average, a one-standard deviation increase in gender equality in one's country of origin (corresponding for example to the difference between the average German-born couple and the average Swedish-born couple according to WVS1) raises couples' response to a cut in the wife's tax rate by about 0.25, and lowers their response to a cut in the husband's tax rate by twice as much, about 0.5.

It is reasonable to expect that, while country-specific indicators capture differences in gender equality and norms pertaining to diverse domains, they are driven by a common, latent determinant at the country level. To explore this possibility we perform a principal-component analysis in column 7 of Table 5. The first principal component (PC) explains about 90% of the overall variance, with an eigenvalue of 5.38. All estimates in column 7 are very similar to those reported in columns 1-6, strongly supporting the hypothesis that the first principal component of these indicators captures the bulk of the role of gender norm variation across couples.

The main takeaway point is that more egalitarian gender norms induce couples to more strongly respond to economic incentives that push towards an egalitarian division of labor, while withstanding incentives that push towards a traditional division of labor. Both effects are statistically significant and quantitatively relevant.

6.3 Robustness

We finally perform some robustness analysis on the heterogeneous effects presented above. First, as spouses in 11% of couples in the sample originate from different countries (up to 17% among immigrants), we test whether norms in the husband's country of origin matter any differently from norms in the wife's country of origin. In Tables C–3 and C–4 of the Appendix we presents results in which norms are measured in the husband's and wife's country of origin, respectively (as opposed to using their average as in Table 5). The estimates obtained are indeed very similar in the two cases. While the proportion of mixed-origin couples is small in our sample, and we are likely underpowered to detect significant differences in behavior, we find no evidence that our results may be disproportionately driven by one spouse's norms as opposed to the other.

Second, we estimate specifications that additionally control for the occupation of each spouse. We noted above that occupation may be an important margin of variation in so far as it proxies work-related constraints in the take-up of TPL, related for example to differential substitutability of occupational roles in the workplace. This may be relevant for our estimates if men and women tend to specialize in different occupations, and differentially so by country of origin. We obtain information on occupation from the Salary Structure Statistics provided by Statistics Sweden, which cover the whole public sector, all private firms with 500 employees or more, and a stratified sample of firms with less than 500 employees. Due to sampling, we can match only about 30% of our original sample to records in the Salary Structure Statistics. Table C–5 in the Appendix reports estimates from specifications that control for 2-digit occupation. The results are closely in line with those reported in Table 5:

all estimates on the interaction terms shown in Panel B remain qualitatively robust, and if anything the point estimates tend to be higher than in Table 5. However, given the reduced sample size, the coefficients of interest are less precisely estimated.

Third, we show estimates in which the dependent variable is residualized with respect to the counterfactual evolution of TPL using the matching, as opposed to the Gelber (2014), method. The results reported in Tables C–6 and C–7 in the Appendix closely replicate the corresponding results based on the Gelber (2014) method in Tables 4 and 5. Our findings are therefore robust to the choice of residualization method.

7 Conclusions

This paper proposes a test of gender identity norms based on the response of husbands' and wives' childcare time to changes in their post-tax wages, which alter the cost of abiding to gendered norms in the division of household tasks.

Based on a household optimization problem, we relate gender norms to the elasticity of substitution between spousal inputs in home production, and argue that asymmetries in such elasticity following cuts in husbands' and wives' tax rates are informative about specific norms – traditional or egalitarian – that a couple abides to. We bring this conceptual framework to the data, combining variation in post-tax wages generated by the Swedish EITC with administrative information on parents' childcare time within the TPL scheme. Our empirical setting allows us to identify the elasticity of substitution between parental inputs in childcare, distinguishing between cases of husbands' and wives' treatment.

We estimate an overall elasticity of substitution of about 1.6, and find evidence of systematic variation in elasticity across couples with different cultural backgrounds. Specifically, couples originating from countries with relatively conservative norms more intensively reallocate childcare across spouses following a reduction in the husband's tax rate, and less intensively following a reduction in the wife's tax rate. These results imply that couples with a more conservative background are more likely to exacerbate gender disparities in childcare time when incentives push in that direction, while they are not as responsive to incentives that would induce a more equal gender division of labor.

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Table 1: Summary statistics

		Mean	SD	Min	Max
Males:	Age	41.1	5.3	21	72
	Education (%)	40.6	49.1	0	1
	Immigrant (%)	11.1	31.4	0	1
	Labor earnings	386	276	39.7	2.95e+04
	Benefits	9.3	20.6	0	382
	Marginal tax rate (%)	43.8	10.8	23.1	59.2
	Days of TPL	2.4	4.6	0	262
Females:	Age	38.7	4.7	19	59
	Education (%)	49.0	50.0	0	1
	Immigrant (%)	11.3	31.7	0	1
	Labor earnings	249	137	39.7	1.19e+04
	Benefit payments	15	27	0	351
	Marginal tax rate (%)	35.4	8.5	23.1	58.7
	Days of TPL	4.9	7.0	0	259
Couples:	Nr. of children aged 4–10	1.5	0.6	1	5
	Age of youngest child	6.8	2.0	4	10
	Male-female gap in taxes (%)	8.4	12.3	-31.7	31.9
	Combined days of TPL	7.3	9.5	0	468
	Male-female gap in TPL	-2.5	7.1	-257	175
Share couples with:	$\text{TPL}_m + \text{TPL}_f = 0$	0.22			
	$\text{TPL}_m = \text{TPL}_f > 0$	0.05			
	$\text{TPL}_m < \text{TPL}_f$	0.54			
	$\text{TPL}_m > \text{TPL}_f$	0.20			
Observations	2006	148,908			
	2007	157,928			
	2008	161,697			
	Total	468,533			

Notes: The table summarizes couples' characteristics as of year $t-1$. All monetary values are expressed in thousand SEK. "Education" takes value 1 if an individual has two years or more of post-secondary education, 0 otherwise. "Benefit payments" include SPL payments, TPL payments, sickness benefits, care allowance, training allowance, unemployment benefits, and rehabilitation compensation.

Table 2: Elasticity of substitution in the main and placebo samples

	Main sample 2006-07, 2007-08 and 2008-09 cohorts			Placebo sample 2004-05 and 2005-06 cohorts		
	Raw (1)	Gelber (2014) (2)	Matching (3)	Raw (4)	Gelber (2014) (5)	Matching (6)
Panel A						
β	2.192***	1.635***	1.708***	0.872***	0.001	0.020
	(0.087)	(0.087)	(0.090)	(0.064)	(0.063)	(0.066)
Panel B						
β^+ $(\tau_f \downarrow)$	1.993***	1.549***	1.696***	0.773***	0.001	-0.011
	(0.110)	(0.110)	(0.114)	(0.081)	(0.084)	(0.084)
β^- $(\tau_m \downarrow)$	2.656***	1.836***	1.738***	1.101***	0.002	0.089
	(0.207)	(0.207)	(0.215)	(0.152)	(0.152)	(0.158)
Row difference	-0.663***	-0.288	-0.042	-0.328	-0.001	-0.100
	(0.256)	(0.256)	(0.265)	(0.189)	(0.189)	(0.196)
N	468,533	468,533	466,420	295,567	295,567	294,869

Notes: The dependent variable in column 1 is the raw change in the log TPL gap between $t-1$ and t (specification (15)); in column 2 it is the residualized change in the log TPL gap based on the Gelber (2014) method and in column 3 it is the residualized change in the log TPL gap based on the matching method (specification (18)). Columns 4-6 report corresponding placebo specifications. Panel B estimates specification (19) with the same dependent variables and samples as in Panel A. All regressions also control for cohort fixed-effects; virtual income, age fixed effects, educational fixed-effects (7 categories), dummy for born in Nordic country, industry fixed effects (10 categories) for each spouse; municipality fixed-effects (289); fixed effects for the number of children aged 4-11; fixed effects for the age of the youngest child; total days of SPL taken by the couple; share of SPL taken by the mother; fixed effects for the number of grandparents living in the same municipality; share of mothers with children aged 0-11 at each spouse's workplace (and their interaction); marital status. Standard errors are clustered at the couple level and reported in brackets. Significance: * = 0.1; ** = 0.05; *** = 0.01.

Table 3: Summary statistics for indicators of gender norms in the working sample

Index	Mean	SD	Min	Max	Obs
WVS1	78.2	13.7	12.3	87.2	468,511
WVS2	74.4	9.1	16.8	91.2	468,511
GGGI	80.1	4.5	55.1	84.5	468,511
SIGI	97.3	6.7	35.9	99.1	468,511
FLFP	85.1	11.1	16.5	89.6	468,511
GDP	54.8	14.1	0.6	96.5	468,511

Notes. Whenever spouses are native of different countries we assign to their couple the average value of the indicator across spouses. All indicators are defined such that higher values correspond to higher gender equality.

- WVS1: % of respondents in the World Value Survey who do not agree with the statement "When job are scarce, men should have more right to a job than women." (waves 5-7).
- WVS2: % of respondents in the World Value Survey who do not agree with the statement "When a mother works, the children suffer." (waves 6-7).
- GGGI: equal to 100 minus the 2016 Global Gender Gap index, capturing gender gaps in economic participation and opportunities, educational attainment, health, survival, and political empowerment; it ranges between 0 "most unequal environment" and 100 "most equal environment" (source: 2016 World Economic Forum's Global Gender Gap Report; 2018 for Iraq).
- SIGI: equal to 100 minus the 2014 Social Institution and Gender Index, capturing discriminatory social institutions, aggregating sub-scores that relate to women's discrimination in the family, restricted physical integrity, restricted access to productive and financial resources, and restricted civil liberties; it ranges between 0 "most discriminatory environment" and 100 "least discriminatory environment" (source: OECD).
- FLFP: ratio of female to male labor force participation rate (x100) in 2011 (source: World Bank).
- GDP: gross domestic product per head in 2011, in thousands USD (source: World Bank).

Table 4: Elasticity of substitution by country of origin

	Birthplace of spouses: Nordic	Non-Nordic	Difference
	(1)	(2)	(3)
Panel A			
β	1.620***	1.835***	-0.215
	(0.089)	(0.341)	(0.351)
Panel B			
β^{+} $(\tau_f \downarrow)$	1.605***	0.630	0.976**
	(0.112)	(0.476)	(0.485)
β^{-} $(\tau_m \downarrow)$	1.627***	3.562***	-1.935***
	(0.218)	(0.626)	(0.659)
Row difference	-0.022	-2.932***	2.910***
	(0.266)	(0.859)	(0.891)
N	434,547	33,986	

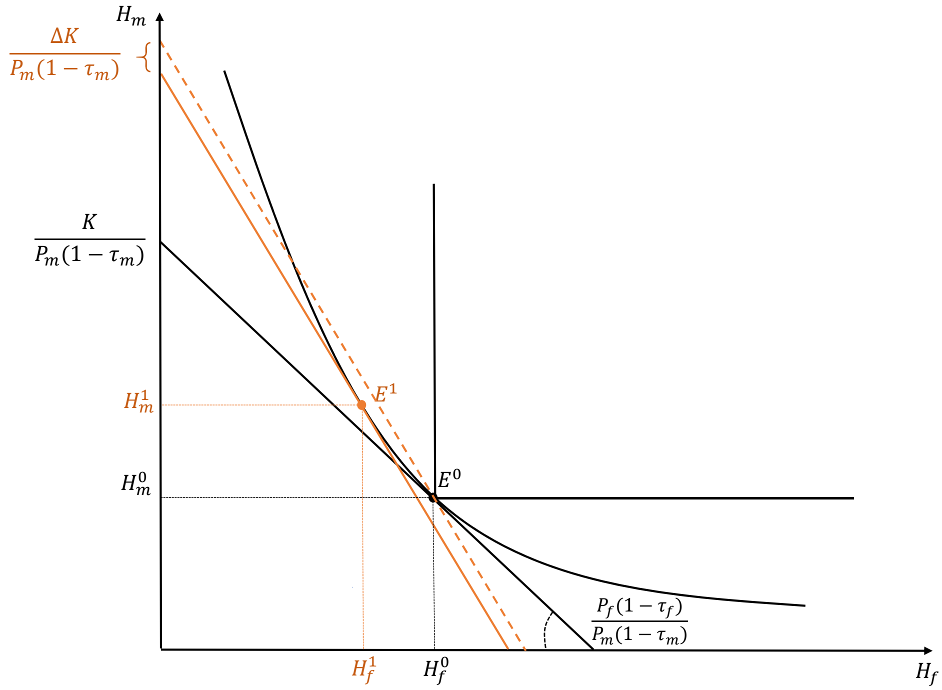
Notes: The dependent variable in all regressions is the residualized change in the log TPL gap based on the Gelber (2014) method. Panel A estimates are based on specification (18) and Panel B estimates are based on specification (19). Coefficients in columns 1 and 2 are estimated in a single regression, including an interaction between the tax variables and a dummy for both spouses being born in a non-Nordic country. Column 3 reports differences between coefficients in columns 1 and 2. Row 4 reports differences between β^{+} and β^{-} estimates for each type of couple in columns 1 and 2, and the corresponding double differences in columns 3. All regressions also control for variables listed in the notes to Table 2. Standard errors are clustered at the couple level and are reported in brackets. Significance: * = 0.1; ** = 0.05; *** = 0.01.

Table 5: Varying elasticities with gender norms in the country of origin

Indicator	WVS1 (1)	WVS2 (2)	GGGI (3)	SIGI (4)	FLFP (5)	GDP (6)	PC (7)
Panel A							
β	1.634***	1.635***	1.634***	1.635***	1.634***	1.635***	1.634***
	(0.087)	(0.087)	(0.087)	(0.087)	(0.087)	(0.087)	(0.087)
$\beta \times indicator$	-0.073	-0.098	-0.068	-0.152*	-0.098	-0.078	-0.098
	(0.090)	(0.091)	(0.089)	(0.087)	(0.089)	(0.091)	(0.090)
Panel B							
β^+	1.534***	1.537***	1.533***	1.546***	1.536***	1.534***	1.536***
	(0.110)	(0.110)	(0.110)	(0.110)	(0.110)	(0.110)	(0.110)
$\beta^+ \times indicator$	0.247**	0.206	0.261**	0.103	0.238*	0.240*	0.231*
	(0.124)	(0.126)	(0.124)	(0.121)	(0.126)	(0.123)	(0.124)
β^-	1.754***	1.763***	1.757***	1.774***	1.756***	1.754***	1.751***
	(0.208)	(0.207)	(0.208)	(0.207)	(0.208)	(0.208)	(0.208)
$\beta^- \times indicator$	-0.542***	-0.533***	-0.539***	-0.517***	-0.554***	-0.568***	-0.560***
	(0.171)	(0.172)	(0.167)	(0.162)	(0.164)	(0.176)	(0.168)
N	468,511	468,511	468,511	468,511	468,511	468,511	468,511

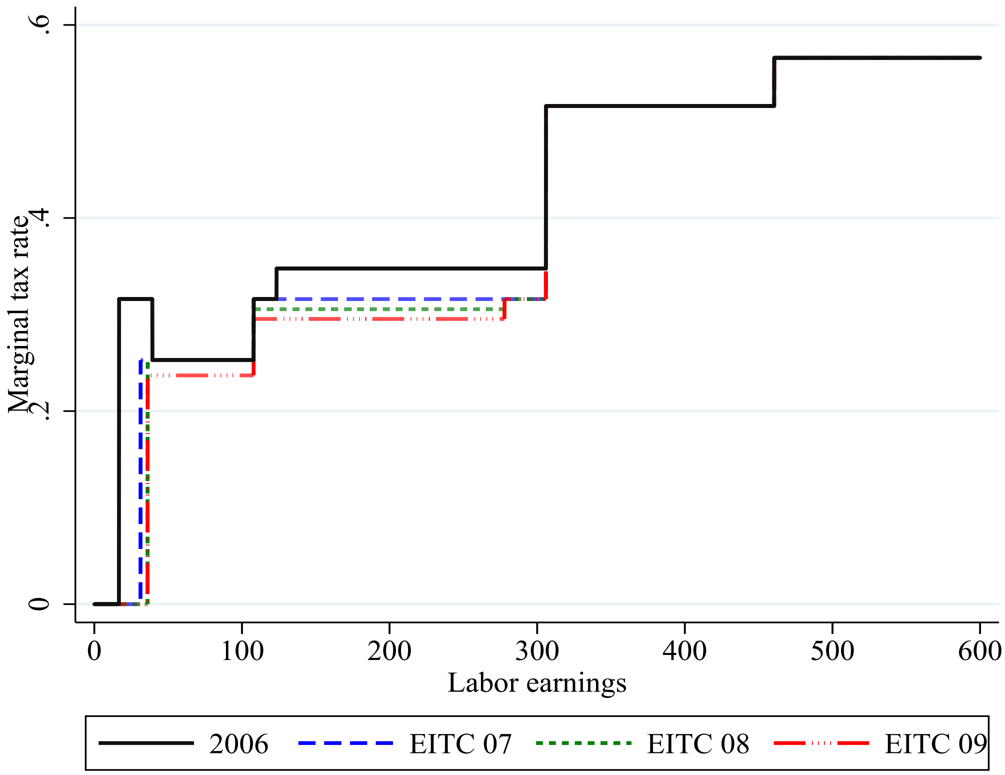
Notes: The dependent variable in all regressions is the residualized change in the log TPL gap based on the Gelber (2014) method. Estimates in Panel A are based on an augmented specification of (18), where $(\tilde{\sigma}_{ift} - \tilde{\sigma}_{imt})$ is interacted with each indicator in turn. Estimates in Panel B are based on an augmented specification (19) where $\Delta(\tilde{\sigma}_{ft} - \tilde{\sigma}_{mt})|_{\Delta\tilde{\sigma}_{ft}>0}$ and $\Delta(\tilde{\sigma}_{ft} - \tilde{\sigma}_{mt})|_{\Delta\tilde{\sigma}_{mt}>0}$ are interacted with each indicator in turn. If spouses are from different countries of origin, we use the average value of each indicator. All indicators are standardized. All regressions also control for the relevant indicator and for variables listed in the notes to Table 2. PC, in Column 7, is the standardized first principal component of indicators 1-6 (explaining 90% of the overall variation, with an eigenvalue of 5.38). Standard errors are clustered at the couple level and reported in brackets. Significance: * = 0.1; ** = 0.05; *** = 0.01.

Figure 1: The impact of a cut in τ_f on couples' time allocation under weak and strong norms



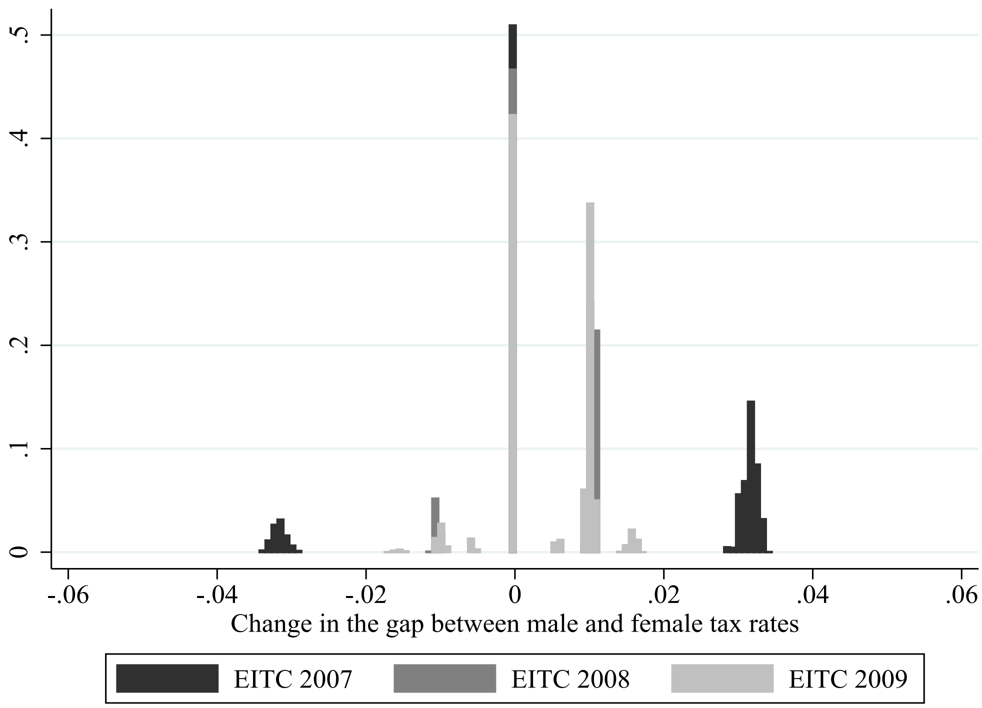
Notes: The figure illustrates the optimal time allocation of two couples, characterized, respectively, by $\beta > 0$ (smooth indifference curve) and $\beta \to 0$ (right-angle indifference curve). Their initial time allocation is identical, and is represented by point E^0. Following a cut in τ_f, the time allocation of the $\beta > 0$ couple moves to position E^1, while the time allocation of the $\beta \to 0$ couple remains unchanged.

Figure 2: Marginal tax rates before and after the EITC's



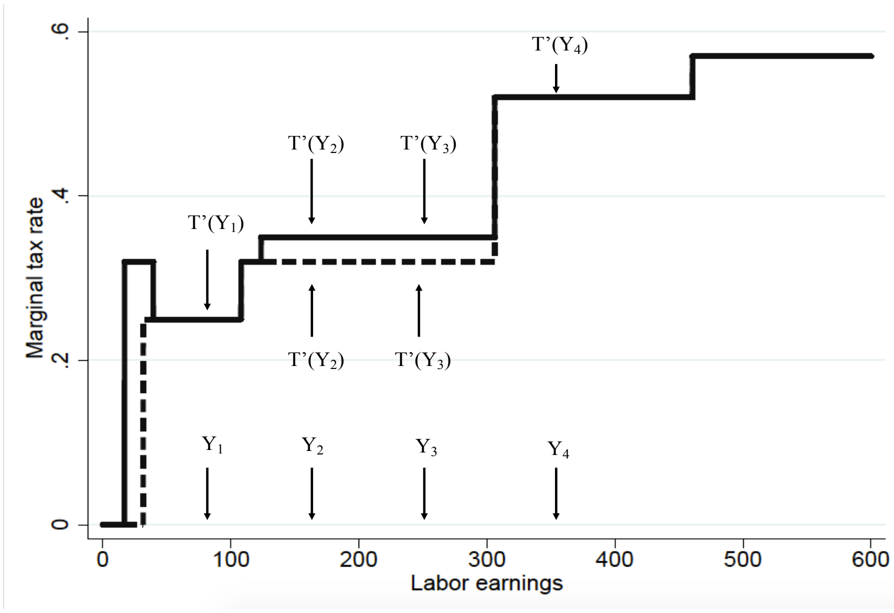
Notes: The solid line represents the tax schedule in 2006. The blue, green and red dashed lines represent changes to the tax schedule introduced with the 2007, 2008 and 2009 EITCs. All marginal taxes are calculated based on a 31.6% municipal tax rate and the 2006 price base amount. Labor earnings are expressed in thousand SEK.

Figure 3: The distribution of simulated changes in marginal tax rates



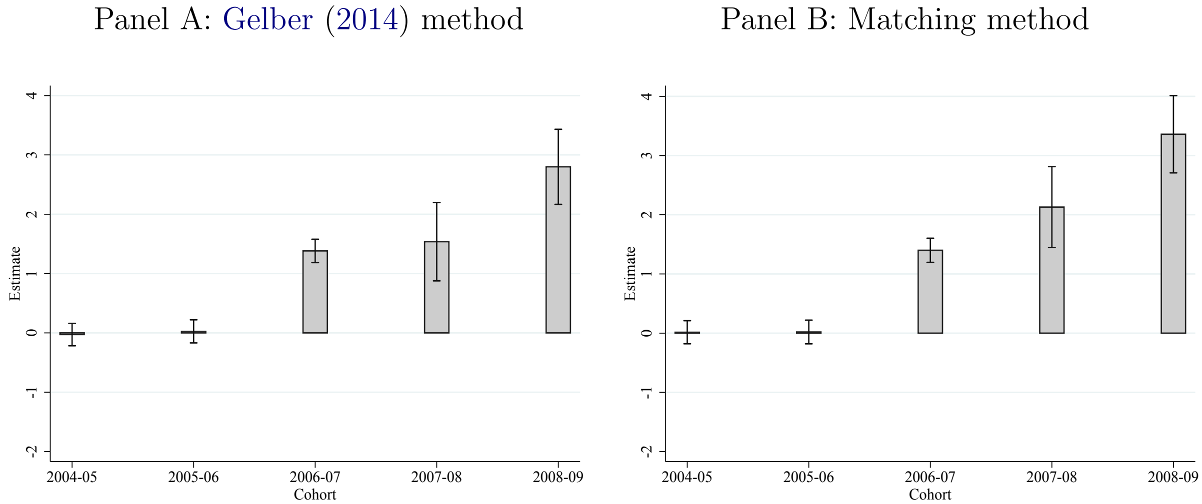
Notes: The histogram represents the distribution of changes in the simulated male-female tax gap induced by the 2007, 2008 and 2009 EITCs.

Figure 4: Examples of control and treated spouses



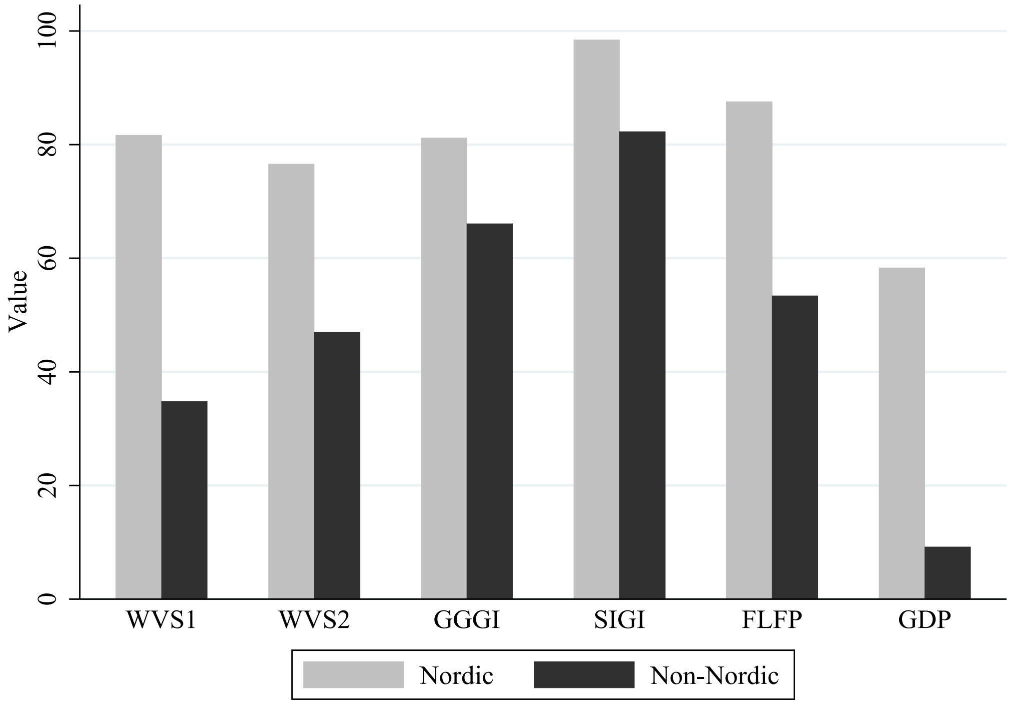
Notes: The solid line represents the tax schedule in 2006. The dashed line shows changes to the tax schedule introduced by the 2007 EITC. All marginal taxes are calculated based on a 31.6% municipal tax rate. The figure indicates earnings and tax rates for four representative individuals of whom: individuals 1 and 4 experience no tax change and individuals 2 and 3 experience identical cuts in marginal tax rates. Labor earnings are expressed in thousand SEK.

Figure 5: Elasticities by year



Notes: The figures plot elasticities separately estimated for five couple cohorts. The 2004-05 and 2005-06 cohorts belong to the placebo sample, which was not exposed to tax reforms. Simulated tax changes for this sample are calculated based on the criteria that determines exposure to the 2007 EITC. The 2006-07, 2007-08 and 2008-09 cohorts constitute our main working sample. The dependent variable in all estimates is the residualized change in the log TPL gap using the Gelber (2014) method in Panel A and the matching method in Panel B. All regressions also control for virtual income, age fixed effects, educational fixed-effects (7 categories), dummy for born in Nordic country, industry fixed effects (10 categories) for each spouse; municipality fixed-effects (289); fixed effects for the number of children aged 4-11; fixed effects for the age of the youngest child; total days of SPL taken by the couple; share of SPL taken by the mother; fixed effects for the number of grandparents living in the same municipality; share of mothers with children aged 0-11 at each spouse's workplace (and their interaction); dummy for marital status. Vertical bars represent 95% confidence intervals using standard errors clustered standard errors at the couple level.

Figure 6: International differences in various indicator of gender norms



Notes: The bars represent alternative indicators of gender norms for Nordic couples (in which both spouses are born in Sweden, Norway, Finland, Denmark or Iceland) and the rest of the sample (Non-Nordic). All indicators are defined such that higher values correspond to higher gender equality. See Table 3 for definitions and details.

Online Appendix

Economic incentives, childcare and gender identity norms

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A The cost of norms: A simple numerical illustration.

Using the standard CES properties for the expenditure function, the cost of achieving utility H when preferences are summarized by an elasticity of substitution β is given by

$$K(\beta, H) = \left\{ \sum_j a_j [P_j(1-\tau_j)]^{1-\beta} \right\}^{\frac{1}{1-\beta}} H, \tag{A–1}$$

with $a_m = s^\beta$ and $a_f = (1-s)^\beta$.

Consider a couple with Leontieff preferences ($\beta \to 0$), an initial home production allocation H_j and associated opportunity cost $\sum_j (1-\tau_j) p_j H_j$, $j = m, f$. Following a change in the tax structure, the home production allocation remains unchanged and its opportunity cost rises to

$$\sum_j (1-\tau'_j) p_j H_j = K'(0, H) = H \sum_j (1-\tau'_j) p_j, \tag{A–2}$$

where τ'_j represents the new tax structure.

If this couple had instead a higher elasticity of substitution $\beta > 0$, the opportunity cost would have risen to $K'(\beta, H) = \{\sum_j a_j [P_j(1-\tau'_j)]^{1-\beta}\}^{\frac{1}{1-\beta}} H$. Given (A–2), $H = K'(0,H)/[\sum_j (1-\tau'_j) p_j]$. The loss of disposable income associated to Leontieff preferences can thus be expressed as a proportion of total home production expenditure $K'(0, H)$:

$$K'(0, H) - K'(\beta, H) = \left\{ 1 - \frac{\{\sum_j a_j [P_j(1-\tau'_j)]^{1-\beta}\}^{\frac{1}{1-\beta}}}{\sum_j (1-\tau'_j) p_j} \right\} K'(0, H). \tag{A–3}$$

We evaluate (A–3) based on total home production time of a representative couple with dependent children in 2006. Based on full-time equivalent (FTE) earnings from the Swedish Wage Structure Statistics and 251 working days in 2006, FTE daily wages in 2006 were equal to $P_m = 1480$ and $P_f = 1160$ SEK. Given the 2007 EITC, these baseline earnings imply $\tau_m = \tau'_m = 0.516$, $\tau_f = 0.347$ and $\tau'_f = 0.316$. We calibrate s using the first order condition (8), together with the above values of P_m, P_f, τ_m, τ_f and our β estimate for the whole sample (1.6, see Table 2). Home production hours are taken from time use surveys. Specifically, Table B:7a in SCB (2012) reports that men and women with children 0-6 do on average 4.6 and 5.7 hours of home production per day, respectively. Corresponding figures for parents of children 7-17 are 3.3 and 4.1, respectively. Using population weights from registry data, we compute that fathers and mothers do on average 4.0 and 4.9 hours of home production per day, corresponding to 1,460 and 1,788.5 annual hours, respectively. These data imply $s = 0.83$. We feed these figures into equation (A–3) to give the loss of disposable income as a function of β. For example, the Leontieff couple would spend an extra $0.24K'(0, H)$ relative to a couple with $\beta = 1$, and an extra $0.34K'(0, H)$ relative to a couple with $\beta = 2$. $K'(0, H)$ represents the daily opportunity cost of home production, as given by (A–2). Using the information described above gives $K'(0, H) = 1,223$ (in 2006 SEK). Thus the Leontieff couple would forgo about SEK 294 per working day with respect to the $\beta = 1$ couple, and SEK 416 with respect to the $\beta = 2$ couple.

B Swedish marginal tax rates

Labor earnings and capital income are taxed at the individual level. For each spouse in our sample we obtain the marginal tax rate by combining tax schedules and annual labor gross earnings. The labor earnings variable in the LOUISE register is the same concept used by Swedish tax authorities to define individual tax liabilities.

The marginal tax rate on labor earnings is a function of a proportional municipality tax, a progressive central government tax, a basic deduction, and an EITC (from 2007). In 2006, the municipality tax rate ranges from 28.9% to 34.2% (with a mean of 31.6%) and the central government tax is phased in at 20% for earnings above 306,000 SEK, rising to 25% for labor earnings above 460,600 SEK.[B–1] Before taxes are levied on gross earnings, a basic deduction is applied, whose value varies with earnings as shown in Table B–1.

Table B–1: Basic deduction 2006-2009

Gross labor earnings (Y)	Basic deduction (BD)
$0 \leq Y < 0.99\times$BA	$0.423\times$BA
$0.99\times$BA $\leq Y < 2.72\times$BA	$0.423+0.20\times(Y$-$0.99\times$BA$)$
$2.72\times$BA $\leq Y < 3.11\times$BA	$0.77\times$BA
$3.11\times$BA $\leq Y < 7.88\times$BA	$0.77\times$BA-$0.10\times(Y$-$3.11\times$BA$)$
$7.88\times$BA $\leq Y$	$0.293\times$BA

Notes: The table shows how to calculate deductions as a function of gross labor earnings. The base amount (BA) is set each year by the Swedish Government to account for inflation and is equal to 39,700 SEK in 2006, 40,300 SEK in 2007, 41,000 SEK in 2008 and 42,800 SEK in 2009.

An EITC on labor earnings was introduced in 2007 and reinforced in 2008 and 2009.[B–2] Table B–2 shows details of the EITC design and Figure 2 illustrates the associated changes in the overall tax schedule.

[B–1] The lower and upper thresholds in 2007/2008/2009 are 316,700/328,800/367,600 SEK and 476,700/495,00/526,200, respectively.

[B–2] With additional reinforcements in 2010, 2014, and 2019, outside our sample period.

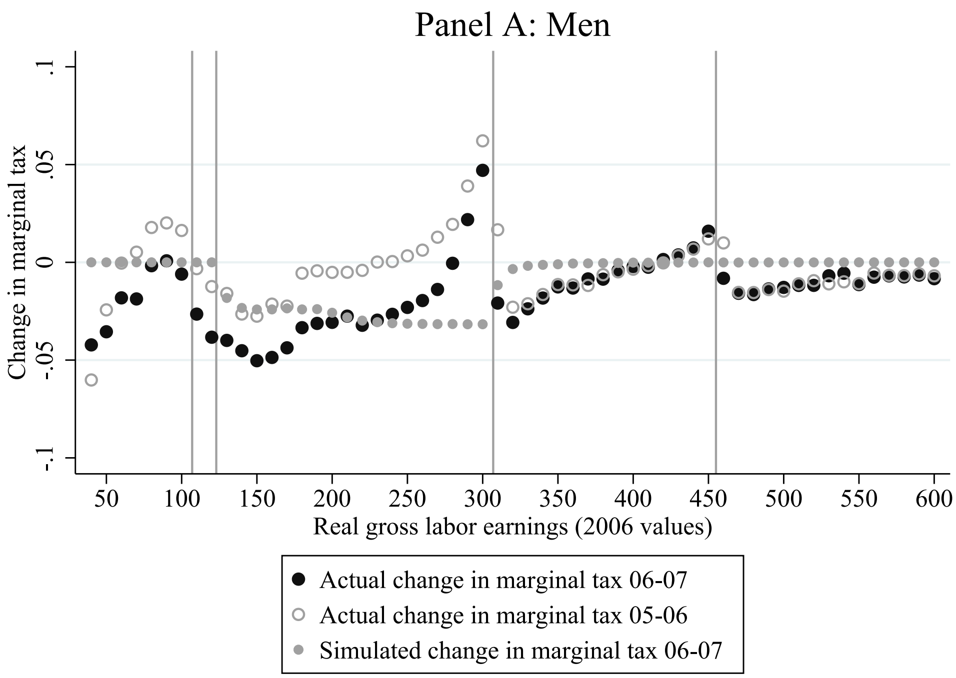
Table B–2: EITC 2007-2009

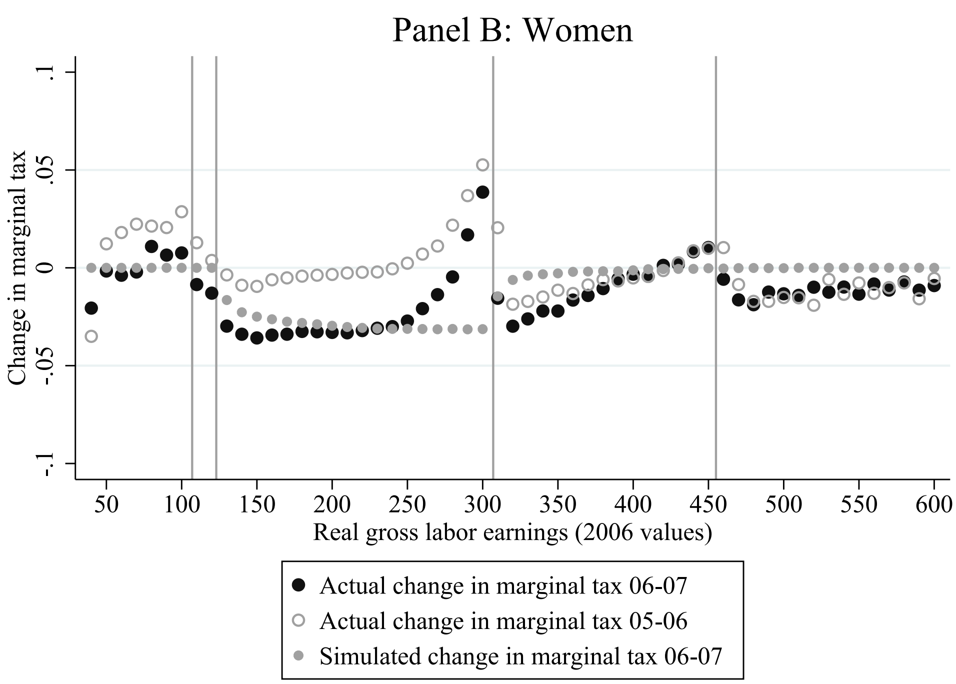
Gross earnings (Y)	EITC
<u>Panel A 2007</u>	
$0 \leq Y < 0.79\times$BA	(Y-BD)×MT
$0.79\times$BA $\leq Y < 2.72\times$BA	(0.79×BA+0.2×(Y-0.79×BA)-BD)×MT
$2.72\times$BA $\leq Y$	(1.176×BA-BD)×MT
<u>Panel B 2008</u>	
$0 \leq Y < 0.91\times$BA	(Y-BD)×MT
$0.91\times$BA $\leq Y < 2.72\times$BA	(0.91×BA+0.2×(Y-0.91×BA)-BD)×MT
$2.72\times$BA $\leq Y < 7\times$BA	(1.272×BA+0.033×(Y-0.72×BA)-BD)×MT
$7\times$BA $\leq Y$	(1.413×BA-BD)×MT
<u>Panel C 2009</u>	
$0 \leq Y < 0.91\times$BA	(Y-BD)×MT
$0.91\times$BA $\leq Y < 2.72\times$BA	(0.91×BA+0.25×(Y-0.91×BA)-BD)×MT
$2.72\times$BA $\leq Y < 7\times$BA	(1.363×BA+0.065×(Y-2.72×BA)-BD)×MT
$7\times$BA $\leq Y$	(1.642×BA-BD)×MT

Notes: The table presents the design of the Swedish EITC introduced in 2007 and further reinforced in 2008 and 2009. BA is the base amount, BD is the basic deduction, MT is the municipal tax rate and Y gross labor earnings.

C Auxiliary tables and figures

Figure C–1: Actual and simulated marginal tax changes





Notes: Changes in marginal tax rates are displayed on the y-axis and real labor earnings (in thousand SEK) are displayed on the x-axis.

Table C–1: TPL use before 2007

Panel A: Base year $t-1$	Combined TPL $H_m + H_f$	TPL gap $H_m - H_f$	log TPL gap $h_m - h_f$
By age of youngest child			
Four	10.52	-3.39	-0.59
Five	9.09	-3.10	-0.58
Six	7.48	-2.61	-0.55
Seven	6.26	-2.27	-0.53
Eight	5.48	-2.07	-0.50
Nine	4.77	-1.82	-0.46
Ten	3.71	-1.51	-0.41

Panel B: Change $t-1$ to t	Change in log TPL $\Delta(h_m - h_f)$
By gender	
(1) Male	-0.06
(2) Female	-0.08
Gap (1)-(2)	0.02
By share of household income	
(1) Main earner	-0.06
(2) Secondary earner	-0.08
Gap (1)-(2)	0.02
By share of TPL	
(1) Main TPL provider	-0.22
(2) Secondary TPL provider	0.08
Gap (1)-(2)	-0.31
N	444,521

Notes: The sample includes the 2004 and 2005 cohorts, selected on the same criteria as cohorts in the main sample (see Section 3.1 for details). $t-1$ and t refers to 2004 and 2005 for the 2004 cohort and to 2005 and 2006 for the 2005 cohort. Panel A displays the average of couples' combined TPL days and the corresponding male-female gap (in levels and logs) by age of youngest child in year $t-1$. Panel B displays the average change from year $t-1$ to t in log TPL by gender, shares of household income and shares of TPL, and the corresponding gaps. Main earners and main TPL providers are defined based on $t-1$ data (when spouses provide identical TPL days, the wife is chosen as the main provider).

Table C–2: Country grouping in registry data

Group	Countries
1	Sweden
2	Finland
3	Denmark
4	Iceland and Norway
5	Bosnia and Herzegovina
6	Croatia, Macedonia, Slovenia and Yugoslavia
7	Poland
8	Ireland and United Kingdom
9	Germany
10	Greece, Italy, Malta, Monaco, Portugal, San Marino and Spain
11	Estonia, Latvia and Lithuania
12	Albania, Armenia, Azerbajdzjan, Belarus, Bulgaria, Georgia, Kazakhstan, Kyrgyzstan, Moldavia, Romania, Russia, Soviet Union, Tajikistan, Turkmenistan, Ukraine and Uzbekistan
13	Czechia, Czechoslovakia, Hungary and Slovakia
14	Andorra, Austria, Belgium, France, Lichtenstein, Luxembourg, Netherlands, and Switzerland
14	Canada and the United States
15	Antigua, Anguilla, Bahamas, Barbados, Barbuda, Belize, Costa Rica, Cuba, Dominica, Dominican Republic, El Salvador, Grenada, Guatemala, Haiti, Honduras, Jamaica, Mexico, Nicaragua, Panama, St Kitts and Nevis, St Lucia and St Vincent
16	Chile
17	Argentina, Bolivia, Brazil, Colombia, Ecuador, Guyana, Paraguay, Peru, Suriname, Uruguay and Venezuela
18	Djibouti, Eritrea, Ethiopia, Somalia and Sudan

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Group	Countries
19	Algeria, Bahrain, Cyprus, Egypt, Gaza, Israel, Jordan, Kuwait, Lebanon, Morocco, Palestine, Qatar, Saudi Arabia, South Yemen, Syria, Tunisia, United Arab Emirates and Yemen
20	Angola, Benin, Botswana, Burkina Faso, Burundi, Cameroon, Cap Verde, Central African Republic, Chad, Comoros, Democratic Republic of Congo, Equatorial Guinea, Gabon, Gambia, Ghana, Guinea, Guinea-Bissau, Ivory Coast, Lesotho, Liberia, Madagascar, Malawi, Mali, Mauritania, Mauritius, Namibia, Niger, Nigeria, Rwanda, Sao Tome and Principe, Seychelles, Sierra Leone, South Africa, Swaziland, Tanzania, Togo, Uganda, Zaire, Zambia, Zanzibar and Zimbabwe
21	Iran
22	Iraq
23	Turkey
24	China, Japan, North Korea and South Korea
25	Burma, Indonesia, Laos, Malaysia, Philippines, Singapore, Thailand and Vietnam
26	Afghanistan, Bangladesh, Bhutan, Brunei, Cambodia, India, Maldives, Mongolia, Oman, Pakistan, Sri Lanka
27	Australia, Fiji, Kiribati, Micronesia, Nauru, New Zealand, Palau, Papua New Guinea, Samoa, Solomon Islands and Vanuatu.

Notes. The table displays the country grouping available in the database LOUISE.

Table C–3: Measuring norms based on husband's origin

Indicator	WVS1 (1)	WVS2 (2)	GGGI (3)	SIGI (4)	FLFP (5)	GDP (6)	PC (7)
Panel A							
β	1.637***	1.638***	1.638***	1.636***	1.636***	1.640***	1.638***
	(0.087)	(0.087)	(0.087)	(0.087)	(0.087)	(0.087)	(0.087)
$\beta \times indicator$	-0.053	-0.083	-0.043	-0.114	-0.085	-0.046	-0.075
	(0.090)	(0.091)	(0.089)	(0.086)	(0.089)	(0.091)	(0.089)
Panel B							
β^+	1.536***	1.541***	1.535***	1.546***	1.538***	1.538***	1.538***
	(0.110)	(0.110)	(0.110)	(0.110)	(0.110)	(0.110)	(0.110)
$\beta^+ \times indicator$	0.255**	0.196	0.275**	0.129	0.231*	0.255**	0.238*
	(0.123)	(0.124)	(0.122)	(0.118)	(0.123)	(0.123)	(0.122)
β^-	1.764***	1.775***	1.770***	1.783***	1.764***	1.769***	1.763***
	(0.208)	(0.207)	(0.208)	(0.207)	(0.208)	(0.208)	(0.208)
$\beta^- \times indicator$	-0.514***	-0.481***	-0.510***	-0.485***	-0.533***	-0.512***	-0.535***
	(0.174)	(0.174)	(0.171)	(0.169)	(0.168)	(0.178)	(0.172)
N	468,511	468,511	468,511	468,511	468,511	468,511	468,511

Notes: This table replicates the specifications of Table 5, having measured each indicator in the husband's country of origin. Standard errors are clustered at the couple level and are reported in brackets. Significance: * = 0.1; ** = 0.05; *** = 0.01.

Table C–4: Measuring norms based on wife's origin

Indicator	WVS1 (1)	WVS2 (2)	GGGI (3)	SIGI (4)	FLFP (5)	GDP (6)	PC (7)
Panel A							
β	1.636***	1.637***	1.636***	1.635***	1.634***	1.638***	1.636***
	(0.087)	(0.087)	(0.087)	(0.087)	(0.087)	(0.087)	(0.087)
$\beta \times indicator$	-0.094	-0.113	-0.094	-0.174*	-0.104	-0.110	-0.120
	(0.090)	(0.090)	(0.089)	(0.087)	(0.090)	(0.090)	(0.090)
Panel B							
β^+	1.541***	1.543***	1.541***	1.549***	1.539***	1.543***	1.542***
	(0.110)	(0.110)	(0.110)	(0.110)	(0.110)	(0.110)	(0.110)
$\beta^+ \times indicator$	0.190	0.167	0.199	0.062	0.216*	0.171	0.181
	(0.122)	(0.123)	(0.123)	(0.122)	(0.127)	(0.121)	(0.123)
β^-	1.766***	1.770***	1.765***	1.778***	1.763***	1.767***	1.759***
	(0.208)	(0.207)	(0.208)	(0.207)	(0.208)	(0.208)	(0.208)
$\beta^- \times indicator$	-0.539***	-0.553***	-0.538***	-0.520***	-0.548***	-0.579***	-0.575***
	(0.171)	(0.176)	(0.169)	(0.163)	(0.166)	(0.179)	(0.171)
N	468,511	468,511	468,511	468,511	468,511	468,511	468,511

Notes: This table replicates the specifications of Table 5, having measured each indicator in the husband's country of origin. Standard errors are clustered at the couple level and are reported in brackets. Significance: * = 0.1; ** = 0.05; *** = 0.01.

Table C–5: Varying elasticities with gender norms in the country of origin
(Controlling for two-digit occupation)

Indicator	WVS1 (1)	WVS2 (2)	GGGI (3)	SIGI (4)	FLFP (5)	GDP (6)	PC (7)
Panel A							
β	1.793***	1.795***	1.795***	1.793***	1.794***	1.794***	1.794***
	(0.171)	(0.171)	(0.171)	(0.171)	(0.171)	(0.171)	(0.171)
$\beta \times indicator$	0.033	-0.061	-0.061	-0.109	-0.118	0.022	-0.052
	(0.171)	(0.175)	(0.172)	(0.151)	(0.171)	(0.172)	(0.169)
Panel B							
β^+	1.645***	1.649***	1.648***	1.667***	1.654***	1.648***	1.651***
	(0.205)	(0.205)	(0.205)	(0.204)	(0.204)	(0.205)	(0.205)
$\beta^+ \times indicator$	0.386*	0.329	0.332	0.093	0.308	0.327	0.311
	(0.216)	(0.220)	(0.217)	(0.196)	(0.219)	(0.215)	(0.215)
β^-	2.065***	2.053***	2.049***	2.097***	2.030***	2.081***	2.049***
	(0.437)	(0.437)	(0.437)	(0.437)	(0.437)	(0.438)	(0.437)
$\beta^- \times indicator$	-0.714**	-0.877**	-0.911**	-0.527*	-1.007***	-0.640*	-0.821**
	(0.357)	(0.370)	(0.358)	(0.310)	(0.349)	(0.372)	(0.352)
N	144,112	144,112	144,112	144,112	144,112	144,112	144,112

Notes: This table reproduced the results in Table 5 when controlling for two-digit occupation fixed effects for each spouse (and their interaction). See Table 5 for details. Standard errors are clustered at the couple level and are reported in brackets. Significance: * = 0.1; ** = 0.05; *** = 0.01.

Table C–6: Elasticity of substitution by country of origin
(Dependent variable residualized using the matching method)

	Birthplace of spouses: Nordic	Non-Nordic	Difference
	(1)	(2)	(3)
Panel A			
β	1.714***	1.633***	0.081
	(0.093)	(0.348)	(0.359)
Panel B			
β^{+} $(\tau_f \downarrow)$	1.755***	0.715	1.040**
	(0.116)	(0.486)	(0.496)
β^{-} $(\tau_m \downarrow)$	1.594***	2.949***	-1.356**
	(0.226)	(0.640)	(0.674)
Row difference	0.161	-2.234***	2.395***
	(0.277)	(0.878)	(0.911)
N	432,472	33,948	

Notes: The dependent variable in all regressions is the residualized change in the log TPL gap based on the matching method. Panel A estimates are based on specification (18) and Panel B estimates are based on specification (19). Coefficients in columns 1 and 2 are estimated in a single regression, including an interaction between the tax variables and a dummy for both spouses being born in a non-Nordic country. Column 3 reports differences between coefficients in columns 1 and 2. Row 4 reports differences between β^{+} and β^{-} estimates for each type of couple in columns 1 and 2, and the corresponding double differences in columns 3. All regressions also control for variables listed in the notes to Table 2. Standard errors are clustered at the couple level and are reported in brackets. Significance: * = 0.1; ** = 0.05; *** = 0.01.

Table C–7: Varying elasticities with gender norms in the country of origin
(Dependent variable residualized using the matching method)

Indicator	WVS1 (1)	WVS2 (2)	GGGI (3)	SIGI (4)	FLFP (5)	GDP (6)	PC (7)
Panel A							
β	1.707***	1.708***	1.708***	1.708***	1.708***	1.708***	1.708***
	(0.090)	(0.090)	(0.090)	(0.090)	(0.090)	(0.090)	(0.090)
$\beta \times indicator$	-0.002	-0.020	0.015	-0.110	-0.024	-0.007	-0.024
	(0.092)	(0.093)	(0.091)	(0.088)	(0.091)	(0.093)	(0.092)
Panel B							
β^+	1.679***	1.682***	1.677***	1.692***	1.681***	1.680***	1.680***
	(0.115)	(0.115)	(0.115)	(0.114)	(0.115)	(0.115)	(0.115)
$\beta^+ \times indicator$	0.260**	0.231*	0.293**	0.125	0.264**	0.246*	0.252**
	(0.127)	(0.129)	(0.127)	(0.124)	(0.129)	(0.126)	(0.127)
β^-	1.678***	1.685***	1.682***	1.682***	1.678***	1.679***	1.674***
	(0.216)	(0.215)	(0.216)	(0.215)	(0.216)	(0.216)	(0.216)
$\beta^- \times indicator$	-0.402**	-0.392**	-0.396**	-0.458***	-0.429**	-0.412**	-0.435***
	(0.175)	(0.176)	(0.171)	(0.165)	(0.168)	(0.180)	(0.172)
N	466,398	466,398	466,398	466,398	466,398	466,398	466,398

Notes: Estimates in Panel A are based on an augmented specification (18) where $(\tilde{\sigma}_{ift} - \tilde{\sigma}_{imt})$ is interacted with the indicators, separately. The dependent variable is the residualized change in the log TPL gap using the matching method. Estimates in Panel B are based on an augmented specification (19) where $\Delta(\tilde{\sigma}_{ft} - \tilde{\sigma}_{mt})|_{\Delta\tilde{\sigma}_{ft}>0}$ and $\Delta(\tilde{\sigma}_{ft} - \tilde{\sigma}_{mt})|_{\Delta\tilde{\sigma}_{mt}>0}$ are interacted with the indicators, separately, with the same dependent variables and samples as in Panel A. All indicators are standardized. PC, in Column 7, is the standardized first principal component of indicators 1-6 with an eigenvalue of 5.38 and the proportion explained by the first principal component equal to 0.90. All regressions control for the relevant indicator and for variables listed in the notes to Table 2. Standard errors are clustered at the couple level and are reported in brackets. Significance: * = 0.1; ** = 0.05; *** = 0.01.