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Essays on Public Policies and Female Labor Supply

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Abstract

In the first chapter of this work, I study the impact of joint taxation of married couples on married couples' labor supply. While joint taxation is fairly widespread across European countries, evidence of its labor supply effects is scarce due to a lack of recent policy changes in family taxation. This chapter makes use of the introduction of joint taxation in the Czech Republic in 2005 to estimate its effect on married couples' labor supply. Results based on difference–in–differences and on triple differences with several alternative control groups suggest that the introduction of joint taxation led to a decline of about 3 percentage points in the employment rate of married women with children. Participation declines are twice as large when the tax work disincentives are highest—among women with tertiary–educated husbands. The introduction of joint taxation did not affect the employment probability of married men with children.

The second chapter contributes to the literature on female labor supply responsiveness by measuring the effect of tax–benefit policies on female labor supply based on a broad sample of 26 European countries in 2005–2010. The tax–benefit microsimulation model EUROMOD is used to calculate a measure of work incentives at the extensive margin—the participation tax rate, which is then used as the main explanatory variable in a female employment equation. This allows me to deal with the endogeneity of income in a new way by using a simulated instrumental variable based on a fixed EU–wide sample of women. Results suggest that a 10 percentage point increase in the participation tax rate decreases the probability of female employment by 2 percentage points. The effect is higher for single mothers, for women in the middle of the skills distribution, and in countries that have lower rates of female employment.

The third chapter explores the effect of two reforms of parental leave allowance in the Czech Republic on the labor market status of mothers with young children. The Czech Republic is a country with a strong attachment of women to the labor market but one of the longest paid parental leave durations. Using a difference–in-differences methodology, I study the effect of two reforms of duration of parental allowance on the labor market status of mothers 2–7 years after childbirth. While the 1995 reform prolonged parental allowance from 3 to 4 years and the 2008 reform introduced a flexible schedule that allowed shortening of leave from 4 to 2 or 3 years, both reforms maintained the job protection period at 3 years, allowing me to study the impact of monetary incentives setting aside changes in job security. I find that the 1995 reform prolonged the parental leave of at least one third of mothers and shifted the post–leave unemployment spell to the time

when a child turns 4, while the 2008 reform achieved a partial reversal of the impact of the 1995 reform.

Abstrakt

V první kapitole této práce zkoumám dopad společného zdanění manželů na nabídku pracovních sil manželských párů. Zatímco společné zdanění manželů je poměrně rozšířené napříč evropskými zeměmi, analýzy jeho dopadů na nabídku práce jsou velmi vzácné kvůli nedostatku reforem v typu zdanění rodiny. Tato kapitola využívá zavedení společného zdanění v České republice v roce 2005 k odhadu jeho dopadu na nabídku práce manželských párů. Výsledky založené na metodě rozdílu–v–rozdílech a metodě rozdílu–v–rozdílech–v–rozdílech s několika alternativními kontrolními skupinami ukazují, že zavedení společného zdanění vede k poklesu míry zaměstnanosti vdaných žen s dětmi o 3 procentní body. Pokles zaměstnanosti je dvakrát tak velký u žen, které zaznamenaly největší pokles pracovních motivací, tedy u žen s vysokoškolsky vzdělanými manželi. Zavedení společného zdanění neovlivnilo pravděpodobnost zaměstnanosti ženatých mužů s dětmi.

Druhá kapitola přispívá k literatuře zabývající se elasticitou nabídky práce žen tím, že zkoumá dopady daňově–dávkových politik na nabídku práce žen na základě mikroekonomických dat z 26 evropských zemí z let 2005 až 2010. Mikrosimulační model EUROMOD je využit k výpočtu ukazatele pracovních motivací—participační daňové sazby, která se pak používá jako hlavní vysvětlující proměnná v participační rovnici. Tento přístup umožňuje vypořádat se s endogenitou příjmů novým způsobem prostřednictvím simulované instrumentální proměnné, která je vytvořena na základě vzorku žen z celé EU. Výsledky naznačují, že nárůst participační daňové sazby o 10 procentních bodů snižuje pravděpodobnost zaměstnanosti žen o 2 procentní body. Efekt je vyšší u svobodných matek, pro ženy se středním vzděláním, a v zemích, které mají nižší míru zaměstnanosti žen.

Třetí kapitola zkoumá vliv dvou reforem rodičovského příspěvku v České republice na ekonomický status matek s malými dětmi. Česká republika je zemí s velmi vysokou mírou zaměstnanosti žen, ale také jednou z nejdelších placených rodičovských dovolených. V této studii používám metodu rozdílu–v–rozdílech pro odhad dopadu dvou reforem rodičovského příspěvku na ekonomický status žen 2–7 let po porodu jejich nejmladšího dítěte. Zatímco reforma z roku 1995 prodloužila délku vyplácení rodičovského příspěvku ze 3 na 4 roky a reforma z roku 2008 zavedla flexibilní systém, který umožnil zkrácení pobírání ze 4 na 2 nebo 3 roky, obě tyto reformy zachovaly délku ochranné doby pro návrat do předchozího zaměstnání na 3 letech, což mi umožnuje oddělit dopady změn peněžních motivací od dopadů změn v délce ochranné doby. Výsledky ukazují, že reforma z roku

1995 prodloužila rodičovskou dovolenou nejméně jedné třetiny matek a posunula období nezaměstnanosti do doby, kdy dítě dosáhne 4 let věku, zatímco reforma z roku 2008 dosáhla opačného efektu, než reforma z roku 1995, ale pouze částečně.

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All errors remaining in this text are the responsibility of the author.

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Klára Kalíšková

Introduction

My dissertation studies the effect of public policies on female work incentives and labor supply decisions. The first chapter examines the impact of family tax treatment on the labor supply of married couples using the introduction of joint taxation in the Czech Republic as a natural experiment. In the second chapter, I study the work incentive effects of tax–transfer policy mix in the EU countries and its impact on the probability of female employment. The third chapter illustrates how two major reforms of parental leave allowance influenced the labor supply of Czech mothers after childbirth over the last 20 years.

The impact of public policies on female labor supply has been widely studied in the economic literature and has brought many important insights over the last few decades. However, there is no clear consensus about the magnitude of labor supply elasticities with respect to various public policies, and the literature is highly concentrated on the US, the UK, and developed economies of Western Europe, while there is little evidence for other countries, including the new EU member states. This thesis contributes to the female labor supply literature by providing new empirical estimates of labor supply elasticities. The first chapter offers new evidence about the magnitude of labor supply elasticities with respect to family tax treatment, which has not been widely studied. The second chapter provides comparable estimates of the labor supply elasticities for 26 EU countries and uses a novel way to deal the endogeneity of income. The third chapter provides new insights into the effect of parental leave policies on the labor supply of women after childbirth from the Czech Republic, a Central European country with very high female labor force participation, but one of the longest paid parental leave durations.

In the first chapter, which was recently published in Labour Economics, I exploit the most recent family taxation reform, the introduction of joint taxation in the Czech Republic in 2005, to estimate its labor supply effects. Even though individual taxation is in force in the majority of EU countries, these countries are not unified in their choice of a tax unit (an individual or a couple). Thus, tax systems based on joint taxation are not exceptional, and tax law often also contains features that provide incentives similar to those of a joint taxation system. However, there is little empirical evidence regarding the labor supply effects of family taxation because of the lack of recent policy changes with respect to family taxation. In this chapter, I apply a difference–in–differences and triple differences approach with several alternative treatment and control groups to evaluate the effect of joint taxation on the married women's and married men's labor supply. The estimates show that joint taxation decreases the labor supply of married women with children—it is associated with a decline of 2.9 percentage points in their employment rate. Moreover, I show that those women who experienced the highest decline in work incentives did indeed respond with the largest decrease in employment probability, and I found no effect on the labor supply of married men.

The second chapter uses an EU–wide tax–benefit microsimulation model EUROMOD to create a measure of work incentives at the extensive margin—the participation tax rate. This measure of work incentives is then used in the female employment equation estimated on a sample of 26 EU countries for 2005–2010. This study makes several contributions to the female labor supply elasticity literature. First, the rich sample allows me to study the responsiveness of the female labor supply across countries and groups of women in a comparable way, and at the same time to control for both time–invariant and time–varying country–level unobserved factors. Second, this paper uses a new approach to deal with the endogeneity of income by using a group–level simulated instrumental variable based on a fixed EU–wide sample of women. My results suggest that a 10 percentage point increase in the participation tax rate decreases the probability of female employment by 2 percentage points. The effect is substantially higher for single mothers, for women with secondary education (as opposed to primary and tertiary educated), and in countries that have lower rates of female employment, such as the countries of Southern Europe.

In the third chapter, which is a joint work with Alena Bičáková, we study the impact of duration of paid parental leave on the post–birth career interruptions of mothers using two reforms of Czech parental leave allowance in 1995 and 2008. The 1995 reform increased the duration of paid parental leave from three to four years. The 2008 reform reduced it for some women to two or three years and made the duration partly choice–based. Although the two reforms altered the duration of the allowance, job protection regulations that entitle women to return to their pre–childbirth jobs remained unchanged, allowing us to study the effect of monetary incentives, setting aside the role of job protection. Despite extensive evidence showing that the initial period of a post–childbirth leave is often followed by a spell of unemployment, previous research focuses solely on the duration of the overall career break. This paper thus also contributes to the existing literature by disentangling the impact of the duration of parental allowance on the overall career break into the initial phase of parental leave (inactivity) and a subsequent period of job search (unemployment). The results suggest that the 1995 reform prolonged parental leave by one year for more than one third of mothers and shifted the occurrence of the spell of unemployment to the time when a child is one year older (4 instead of 3). The 2008 reform, on the other hand, shortened the parental leave of at least one fifth of mothers and shifted the post–leave job search to the time when a child turns 2 or 3.

Chapter 1

Labor Supply Consequences of Family Taxation: Evidence from the Czech Republic[1]

1.1 Introduction

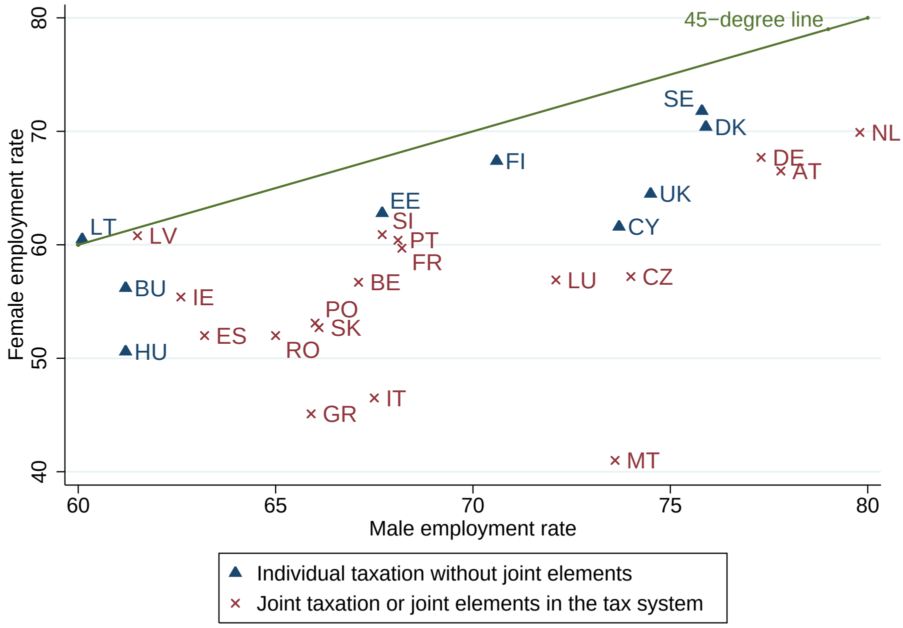
The choice of an appropriate family tax treatment is an important part of an optimal tax design. While individual taxation systems tax each individual's income separately, systems of joint taxation either tax the sum of the family income as a whole or tax each spouse individually based on half of the total income (Stephens and Ward-Batts 2004). Joint taxation meets the requirement for equal treatment of households with the same total income—the tax liability of a married couple is the same regardless of how income is divided between spouses (Cigno, Pestieau, and Rees 2011). However, joint taxation equalizes the marginal tax rates of the spouses and thereby decreases the marginal tax rates of primary earners (usually men) and increases the marginal tax rates of secondary earners (usually women). The effect of joint taxation on the labor supply of married men is ambiguous, because the substitutions and income effects work in opposite directions, but the theoretical effect on the labor supply of married women is unambiguously negative. This study is concerned with the empirical investigation of these labor supply aspects of family tax treatment.

Countries are not unified in their choice of tax unit. Even though individual taxation is in force in the majority of EU countries, a tax law often contains features that provide

[1] **This work has been published in Kalíšková, K. (2014). Labor Supply Consequences of Family Taxation: Evidence from the Czech Republic. *Labour Economics*, 30, 234-244.**

incentives similar to those of a joint taxation system, and tax systems based on joint taxation are not exceptional either.[2] Figure 1.1 shows that, indeed, countries with systems of (truly) individual taxation (Sweden, Denmark, Finland, the United Kingdom, etc.) tend to have higher female employment rates (for a given level of male employment) than countries with joint taxation systems or systems with 'joint' features.

Figure 1.1: Female and male employment rates by taxation systems in the EU.



Note: The graph shows employment rates (15 to 64 years) in 2011. Joint elements in the individual taxation systems are tax deductions for single–earner couples. Individual taxation systems without joint elements: BU, CY, DK, EE, FI, HU, LT, SE, UK; individual taxation systems with joint elements: AT, CZ, EL, IT, LV, NL, RO, SI, SK; and joint taxation countries: BE, DE, ES, FR, IE, LU, MT, PO, PT. Source: Eurostat LFS employment statistics and EUROMOD country reports 2007–2010: `http://epp.eurostat.ec.europa.eu/portal/page/portal/statistics`; `https://www.iser.essex.ac.uk/euromod/resources-for-euromod-users/country-reports`

Although economic theory predicts a negative female labor supply effect of joint taxation, there is little empirical evidence as a result of the lack of recent policy changes with respect to family taxation. Two studies have estimated the impact of joint taxation on the labor supply of married women using family taxation reforms: LaLumia (2008) and Selin (2009). Although both studies provide a comprehensive analysis of the changes in the tax treatment of families in the U.S. and Sweden respectively, their results are based on tax reforms that are more than 40 years old. Among others, Blau and Kahn (2007)

[2] Among others, Crossley and Jeon (2006) argue that 'joint' elements in the individual taxation systems (mainly tax deductions for single–earner couples) provide incentives similar to joint taxation. About one third of EU countries have individual taxation systems with these 'joint' elements, and about one third have joint taxation systems (see note below Figure 1.1).

show that the female labor supply elasticities and behavioral responses to tax reforms have changed significantly since the 1980s, pointing to the need for more up–to–date evidence.

This paper exploits the most recent family taxation reform, the introduction of joint taxation in the Czech Republic in 2005, to estimate the labor supply effect of joint taxation.[3] From January 1, 2005, married couples raising at least one child could have taken the opportunity for joint taxation in the Czech Republic. Since the actual usage of joint taxation among eligible couples is unknown, what I estimate here is the intention–to–treat effect of this reform.[4]

I apply a difference–in–differences approach with several alternative treatment and control groups to evaluate the effect of joint taxation on the married women's and married men's labor supply. The whole analysis is conducted separately by gender. First, I compare married individuals with children (all eligible) with unmarried individuals and married individuals without children (all ineligible). Next, I use the discontinuity in the eligibility rule—children are defined by a strict age threshold in the Czech tax code, which is 18 years, or 26 years in the case of full–time students. Therefore, I focus on a more homogeneous subset of the sample and compare married individuals with children aged 10–17/25 and married individuals with children aged 18/26–30. Furthermore, I apply a local difference–in–differences estimation around the two age thresholds—comparing married individuals with children aged 16–17 vs. those aged 18–19 (not in education), and married individuals with children aged 24–25 (in education) vs. those aged 26–27. Finally, I provide several robustness checks including the triple differences estimation (with an additional control group of Slovak married individuals with children)[5] and two placebo tests to check the validity of the estimation approach.

This project sheds new light on the effect of the family tax treatment on the labor supply of married men and women with children. The estimates show that joint taxation decreases the labor supply of married women with children—it is associated with a decline of 2.9 percentage points in their employment rate. Moreover, I show that those women

[3]The second most recent tax reform concerning family taxation was in the UK in 1990 (the abolition of joint taxation).

[4]The voluntary nature of joint taxation is not uncommon in the European tax systems. It is actually used in the majority of countries that have joint taxation (Germany, Ireland, Malta, Poland, Portugal, and Spain). The intention–to–treat might thus be the main parameter of interest for policy makers. Nevertheless, I show in Section 1.3 that the intention–to–treat effect estimated in this study (the effect of having the joint filing option) provides a lower bound for the effect of mandatory joint taxation.

[5]This is motivated by a common history of the Czech and Slovak Republics and by the fact that labor supply decisions in these countries have many common features even today (Bičáková 2010).

who experienced the highest decline in work incentives did indeed respond with the largest decrease in employment probability (by 5.5 percentage points).

The estimated effect for married men is largely insignificant at the extensive margin, supporting the findings of LaLumia (2008), who also did not find any effect of joint taxation on the labor supply of married men. The effect of joint taxation on hours worked by married men is negative and significant in most specifications, but of a very small magnitude, which is consistent with the income effect slightly outweighing the substitution effect.

The remainder of the paper is organised as follows. The next section reviews the relevant literature, then the institutional background of the Czech reform analysed in this paper is introduced, with an ensuing discussion of the methodology and identifying assumptions of the chosen approach. Finally, the paper presents the results and concludes.

1.2 Literature review

Recently, there has been an expansion in the literature that simulates the effect of a switch from joint to individual taxation on female labor supply (among others, see Steiner and Wrohlich 2004; or Haan 2010). However, these microsimulation studies face common problems connected to the estimation of labor supply effects. Blundell, Duncan, and Meghir (1998) argue that the "[l]abor supply effects have been notoriously difficult to estimate in a robust and generally accepted way" (p. 827). The main reason is the presence of severe simultaneity problems with wages and other income. However, Blundell, Duncan, and Meghir (1998) point out that these estimation problems can be solved if researchers correctly exploit the variation induced by tax reforms. Tax reforms provide us with an exogenous variation in the after–tax wages and enable the observation of behavioral responses to the tax reforms.

This study is highly motivated by these considerations, and I thus base my analysis on the actual policy change. To my knowledge, there are only two studies that use policy reforms in estimating the labor supply effect of joint taxation, and they are based on tax reforms that are more than 40 years old.[6] LaLumia (2008) uses the difference–in–differences strategy at the state level taking advantage of the U.S. tax reform which

[6] There is a related literature focusing on the labor supply effects of more recent tax reforms that introduced flat taxation in Russia and some European countries (see e.g. Duncan and Sabirianova Peter 2010).

introduced joint taxation in 1948. Selin (2009) studies the abolition of joint taxation in Sweden in 1971. Both studies have found a significant impact of family taxation policies on female labor supply decisions, but of different magnitudes. LaLumia (2008) found the effect only among women in highly–educated couples, and of a lower magnitude (2 p.p. decrease in the employment rate) than Selin (2009), who estimates the effect to be about a 10 p.p. increase in the employment of married women. Further, LaLumia (2008) finds that joint taxation did not affect the labor supply of married men. This paper contributes to this literature by providing up–to–date evidence on the impact of joint taxation on the labor supply of married men and women with children. Since the female labor supply elasticities have changed substantially and the amount and composition of women participating in the labor market has also changed drastically over the last 40 years (see e.g. Blau and Kahn 2007), the estimated effects of joint taxation might be very different nowadays compared to four decades ago.

1.3 Institutional background

The policy change of interest in this study is the introduction of joint taxation in the Czech Republic in 2005. Before 2005, there was a progressive individual tax system with four tax brackets.[7] Married couples were taxed individually based on each individual's income. Joint taxation was in force from January 1, 2005, and did not change the structure of the tax brackets, but allowed married couples with children to be taxed based on half of their total income. Joint filing was voluntary, and this option was given to married couples raising at least one child (throughout the paper, I define children consistently with the tax law as those under 18 or 26 in the case of full–time students). Therefore, married couples with children could either choose to be taxed jointly based on half of the total income or individually based on each spouse's income.

In 2008, joint taxation was abolished in the Czech Republic, because flat tax was introduced. However, the effect of the abolition of joint taxation cannot be separated from the effect of the flat tax reform, because the latter was accompanied by an extremely large increase in the tax deduction for single–earner couples that significantly decreased the work incentives of married women.[8] For this reason, I concentrate solely on the

[7] The tax rates were 15% for the tax base below CZK 109,200; 20% for the tax base between CZK 109,200 and 218,400; 25% for the tax base between CZK 218,400 and 331,200; and 32% for the tax base above CZK 331,200.

[8] Tax deduction for single–earner couples was increased from CZK 350 to CZK 2,070 monthly. There-

impact of the introduction of joint taxation.

While joint filing was voluntary, this option was widely used. The official statistics of the Czech Ministry of Finance report that 32.3% of all tax returns in 2005, 35.7% in 2006, and 40.3% in 2007, were filed jointly, while the approximate share of the working population eligible for joint taxation was close to 47% in all relevant years.[9] Although the estimated usage of joint taxation is quite high (69% in 2005, 76% in 2006, and 86% in 2007), I next discuss the implications of the fact that joint filing was voluntary.

Since the data used in the estimation has no information about the actual usage of joint taxation, this paper estimates the effect of the introduction of joint taxation that should be interpreted as the effect of an intention to treat (ITT), i.e. the average causal effect of assignment on the outcome (see e.g. Angrist, Imbens, and Rubin 1996). The assignment here is the opportunity to file taxes jointly and the outcome is the labor supply of married men and women with children. Although the selection to treatment (the actual usage of joint taxation) is likely to be non–random, the assignment to treatment is defined strictly based on eligibility conditions and there is no voluntary component to it. Therefore, in what follows I refer to the assignment to treatment (the option to file jointly) as the treatment.

The fact that joint taxation was voluntary implies that it was chosen only by married couples for whom it lowered the taxes when compared to individual taxation. In general, married couples can be divided into three groups based on how the introduction of joint taxation affects their tax duty. While most of the couples benefit from joint taxation in terms of lowering their tax duty, there are some couples for whom the type of the taxation system does not matter, and some couples who would be better off under individual filing. First, joint taxation is beneficial (in terms of decreasing their tax duty) for all one–earner couples, but also for those two–earner couples who have a sufficiently unequal distribution of labor income between the spouses (approximately 76.6% of married couples with children[10]). Second, the type of family taxation does not matter for two–earner couples where both spouses earn income that falls in the same tax bracket (approximately 20% of married couples with children). Third, the group of couples who are better off

fore, a husband paid CZK 1,720 less on taxes per month if his wife was not working.

[9] The share of eligible couples was calculated based on the Czech Labor Force Survey data 2005–2007. Married couples with at least one child and at least one of the spouses working were considered eligible for joint filing.

[10] The share of couples that belong to each group was calculated using the Czech SILC (Statistics on Income and Living Conditions) data for 2005. SILC is being collected annually by the Czech Statistical Office as a part of the EU SILC project.

under individual taxation includes two–earner couples with one spouse earning income that falls into a tax bracket that is immediately above the tax bracket of the other spouse, and with one of the spouses' income being close to the upper bound of his/her tax bracket (approximately 3.4% of married couples with children). For this small group of couples, the combined tax duty falls in the tax bracket of the higher–earning spouse causing the tax duty under joint taxation to be higher than under individual taxation.

Therefore, the voluntary nature of joint taxation only reduced the difference between tax duty paid in a situation of both spouses working and a situation of only one spouse working for this small groups of couples. In other words, the negative effects of joint taxation on the secondary earner's work incentives were slightly diminished by the fact that joint taxation had a voluntary component. Therefore, the estimated effect of joint taxation in this study should be considered a lower bound of the effect for a similar reform without the voluntary component. However, the difference in the estimated effect is likely to be small, because the share of couples for whom the voluntary component mattered was very low (only about 3.4% of married couples with children, see above).

The Czech Republic is a country with high labor force participation rates and relatively small labor supply elasticities compared to other EU countries (Bičáková, Slačálek, and Slavík 2011). The estimated effects are thus likely to provide a lower bound for the effect of joint taxation than could be expected in other EU countries. Furthermore, the availability of jobs with other than standard full–time working hours is very low in the Czech Republic (see e.g. Tang and Cousins 2005). Therefore, the intensive margin effect could also be much higher in countries with higher labor market flexibility.

I turn now to illustrate the magnitude of joint taxation impact on work incentives. In what follows, it is assumed that men are primary earners and women are secondary earners. This is largely confirmed by the data, as 84% of Czech married women earn less than their husband.[11] The theoretical impact of joint taxation on the labor supply of primary earners (men) is ambiguous, because the income and substitution effects work in opposite directions.[12] The theoretical effect on the labor supply of secondary earners (women) is unambiguously negative, but the magnitude of the impact differs substantially across groups of women by their and their husband's income. Table 1.1 illustrates this

[11] This is calculated using the Czech SILC data for years 2004 to 2007.

[12] For primary earners, the substitution effect increases work incentives, because the tax rate decreases, but the income effect decreases work incentives, because of an increase in family income. For secondary earners, both the substitution and the income effects decrease work incentives, because of an increase in the tax rate and in family income.

heterogenous impact of the introduction of joint taxation on female work incentives by showing the net gain from a wife's work (difference between family income if the wife works and if she does not work) for various combinations of tax brackets of the wife and the husband.[13]

Table 1.1: Work incentive effects of the introduction of joint taxation in the Czech Republic

Tax bracket husband	Tax bracket wife	Change in net gain from a wife's work as a result of the introduction of joint taxation	
		in CZK per month	as % of wife's gross wage
1	1	-257	-2.7%
2	1	-384	-4.1%
2	2	-461	-2.8%
3	1	-1,552	-16.5%
3	2	-1,302	-7.9%
3	3	-2,312	-9.1%
4	1	-1,162	-12.4%
4	2	-5,190	-31.5%
4	3	-5,176	-20.4%
4	4	-3,846	-7.6%

Note: The net gain from a wife's work is the difference between the net household income when the wife works and when she does not work. It is calculated for the average male/female wage in each tax bracket in 2004 (taken from SILC data) and for a family with two children. Calculations take into account not only the effect of income taxes, but also the effect of social benefits, and they are based on the Czech legislation as of 2004. I assume full tax compliance and a full take–up of social benefits. There was a progressive income tax system with four tax brackets (the tax rates were 15% for the tax base below CZK 109,200; 20% for the tax base between CZK 109,200 and 218,400; 25% for the tax base between CZK 218,400 and 331,200; and 32% for the tax base above CZK 331,200) in 2004.

Clearly, there was a substantial decrease in work incentives for wives in the joint (as opposed to individual) taxation system as illustrated by a decrease in the net gain from work in Table 1.1. Also, the magnitude of the disincentive effect of joint taxation increased substantially with the husband's tax bracket—while women whose husbands earned incomes belong in the first tax bracket experienced only a very small negative impact on their work incentives (decrease in net gain from work of CZK 257 per month

[13]Table 1.1 reports the net gain from a wife's work for women who are secondary earners, i.e. their tax bracket is lower or equal to their husband's tax bracket. For details on the calculation, see note below Table 1.1.

or 2.7% of wife's gross wage), the effect was 5 to 10 times higher if the husbands' income belonged in the third tax bracket, and 15 to 20 times higher if it belonged in the fourth tax bracket in absolute terms.

In the period of interest, other tax and benefit reforms increased support for families with children. First, the amount of birth grant was doubled in 2006 and parental allowance benefit was substantially increased in 2007.[14] While these reforms could have affected the fertility decisions (see Section 1.4.3 for discussion), it is unlikely that this happened within the time period in question. Further, these reforms did not affect the labor supply of married couples directly, because these benefits are not means–tested. They could have affected the labor supply of married individuals with small children (up to four years of age) through the income effect, but this could jeopardise the validity of only one of the treatment and control groups used in this study (for details, see Section 3.3.1). Further, there were several reforms of child tax allowances that slightly increased support for low–income families.[15] Since these changes were very small compared to the introduction of joint taxation and the child tax allowance can be used by one parent only, the reforms affected the work incentives of secondary earners only through the income effect, and the magnitude of the impact was likely small.

1.4 Methodology and data

1.4.1 Simple model of family labor supply

My empirical strategy is based on a simple model of family labor supply, which is often referred to as a unitary model (Samuelson 1956). This model treats the household as a single decision–making–unit assuming that spouses pool their resources and maximize joint utility.[16] Moreover, following Eissa and Hoynes (2004) and LaLumia (2008), I

[14] The birth grant is a one–off benefit given to any mother who gives birth to one or more children. The amount of birth grant was doubled in 2006, from CZK 8,750 to CZK 17,500 for a single child. The parental allowance is a benefit for parents who take care for a child up to four years old on a daily basis. In 2007, the monthly allowance was increased from CZK 3696 to CZK 7580.

[15] In 2004, the child tax base deduction was increased from CZK 23,520 to 25,560 per year. In 2005, the child tax base deduction was replaced with a child tax credit of CZK 6,000 per year. This reform increased the monthly amount of child tax allowance by CZK 180 for the lowest tax bracket and decreased it by CZK 182 for the highest tax bracket. Compared to joint taxation, which changed the tax duty by as much as CZK 4,000 per month for women with high–income husbands (see Table 1.1), these were very small changes.

[16] An alternative approach to modeling family structure is a collective model of household labor supply (see e.g., Apps and Rees 1999), which is based on individual decisions and assumes that they lie on the Pareto frontier. However, the collective models have so far been of limited use in an empirical analysis

assume that the primary earner makes his work decision independent of the secondary earner, but the secondary earner takes into account the primary earner's decision. This simple model can be summarized by a pair of labor supply equations (Eissa and Hoynes 2004):

$$H_1 = h_1(w_1, Y, X) \text{ and } H_2 = h_2(w_2, Y + w_1 H_1, X), \tag{1.1}$$

where H_1 and H_2 are hours worked by primary and secondary earners at wages w_1 and w_2, respectively; Y is family non–labor income, and X represents family characteristics.

1.4.2 Difference–in–differences and triple differences approach

I base my empirical strategy on a difference–in–differences approach focusing on the family taxation reform in the Czech Republic in 2005, which introduced joint taxation of married couples with children (up to the age threshold for children defined by Czech law, which is 18 or 26 in the case of full–time students). I define control groups based on the eligibility rules for joint taxation and I conduct the analysis separately by gender. A natural starting point of the analysis is to use all ineligible individuals (unmarried and married without children) as a control group. A similar approach has been used for evaluating the labor supply effects of various policy reforms (see e.g. Meyer and Rosenbaum 2001, Eissa and Hoynes 2004, Mogstad and Pronzato 2012). However, the fact that joint taxation was available only to families with children (strictly defined by the age threshold) gives a unique opportunity to study the effect using supposedly more comparable groups of men/women, which differ only by the age of the youngest child in a family. Therefore, I narrow the analysis and compare married individuals with children aged 10 to 17/25 (a subset of treated individuals) with a control group of married individuals with children aged 18/26 to 30.[17]

Next, I further narrow down the definitions of treatment and control groups and focus on married men/women with children who are just below or just above one of the age thresholds defined by the Czech tax code. In particular, I compare married individuals with children aged 16 or 17 with married individuals with children aged 18 or 19 (who are not in education), and married individuals with children aged 24 or 25 (who are in

of changes in a tax law (Eissa and Hoynes 2004).

[17]This control group thus consists of married couples with children who live in the same household, but are no longer perceived as children by the tax code, because they are older than 17 and they are not full–time students or they are full–time students but are older than 25.

education) with married individuals with children aged 26 or 27. Table 1.2 summarizes the treatment and control groups that are used in the difference–in–differences estimation.

Table 1.2: Difference–in–differences: summary of treatment and control groups.

	Treatment group	Control group
1	Married men/women with children (aged 0–17/25)	Unmarried and married men/women without children (or with children aged over 18/26)
2	Married men/women with children aged 10–17/25	Married men/women with children aged 18/26–30
3	Married men/women with children aged 16–17	Married men/women with children aged 18–19 (not in education)
4	Married men/women with children aged 24–25 (in education)	Married men/women with children aged 26–27

For each of the above mentioned treatment and control groups, I estimate the following equation:

$$Y_{it} = X_{it}\theta + \beta\delta_{gt} + \gamma_t + \gamma_g + \epsilon_{it}. \tag{1.2}$$

The outcome of interest (Y_{it}) is the measure of labor supply at the extensive (dummy equal to one if the man/woman was employed last week) and intensive margin (number of hours worked if employed). The equation includes fixed group and fixed time effects (γ_g and γ_t, respectively) to control for time–invariant group differences and for the common time trend in the labor supply. The impact of joint taxation is captured by β, which is the coefficient of the indicator variable for the treated group in 2005–2007, the years when joint taxation was in force. X_{it} represents a set of observable characteristics including age, education dummies, number of children of a certain age (aged 0–2, 3–5, 6–9, 10–14, 15–17), a dummy variable for cohabiting and married individuals, dummies for the education of a spouse, a dummy variable for the spouse being inactive, number of household members, dummy variables for the non–Czech nationality (either EU nationality or non–EU nationality), and regional dummies.

I provide a more detailed analysis of the treatment effect for women by interacting β with characteristics of the husband and the wife (husband's education and husband's and wife's education) to capture the differences in the intensity of treatment across treated women (for details, see Section 1.3).

As a robustness check, I use a triple differences approach with Slovak men/ women serving as a second control group, which is motivated by the fact that the labor supply decisions of Czech and Slovak couples follow similar patterns (see e.g., Bičáková 2010). Slovak married men/women with children cannot be used directly in the difference–in–differences estimation, because Slovakia experienced a major tax reform in 2004 that also affected working incentives of married couples.[18] However, the effect of the Slovak tax reform (as well as all other country–specific policy reforms) can be filtered out in the triple differences approach. Apart from using the control group of Slovak married men/women with children, I use a second group—married men/women with children aged 18/26–30. This second control group faced the same policy changes concerning tax and social systems in a particular country as married men/women with children aged under 18/26,[19] but this group was not affected by joint taxation policies. Therefore, in the triple differences estimation strategy I difference over time (the before/after difference), across states (the Czech/Slovak difference), and across groups of men/women (the difference between the treatment group of married men/women with children aged 10–17/25 and the control group of married men/women with children aged 18/26–30).

The triple differences estimation equation takes the following form:

$$Y_{ict} = X_{ict}\theta + \beta\delta_{gct} + \gamma_{gt} + \gamma_{ct} + \gamma_{gc} + \epsilon_{ict}, \tag{1.3}$$

where Y_{ict} is the outcome variable (employment dummy/hours worked) and X_{ict} is a set of observable characteristics. I also include group–year, country–year, and group–country interaction terms (γ_{gt}, γ_{ct}, and γ_{gc}, respectively) which capture the differences in trends in employment and hours worked across the two countries, across the two groups of men/women, and the differences in tastes for work between the two groups of men/women

[18] In 2004, Slovakia replaced its progressive tax system (with tax rates varying from 10 to 38%) with a flat tax rate of 19%. This tax reform was accompanied by a significant increase in a tax allowance for single–earner couples (from SKK 12,000 to SKK 87,936 per year) substantially decreasing the work incentives of Slovak married women.

[19] The only exception are changes in policies connected to the presence of children in a family (such as a child tax credit). However, since the child tax credit can be used by one of the spouses only, changes in this tax credit do not affect the labor supply of women with working husbands as the husbands use this tax relief. Moreover, although the Slovak flat tax reform changed the child tax base deduction to a child tax credit, the monetary value of this tax relief changed only a little (by SKK 260 per month for the lowest tax bracket and by SKK 132 for the highest).

in the two countries. The effect of joint taxation is captured by β coefficient, which is a coefficient of the indicator variable for the treatment group (married men/women with children aged 10–17/25) in the Czech Republic in 2005–2007.

1.4.3 Identification assumptions

For the difference–in–differences approach to be valid, two identification assumptions need to be satisfied. First, in the absence of any treatment (without changes in family tax policy), the trends in the labor supply of treatment and control groups would have been the same. Similarly, the triple differences approach requires that the group differences (differences in labor supply between the treatment and control groups) follow the same trend in the two countries. The second assumption requires no significant composition changes in the treatment and control groups.

To provide some evidence concerning the first identification assumption, I plot the evolution of employment–to–population ratios and hours worked for the treatment and control groups in Figure 1.2 and Figure 1.3.[20] The two graphs in the upper part of Figure 1.2 compare the employment–to–population ratios for the first two treatment and control groups of women. The trends in employment for the treatment and control groups of women seems to be similar for the early pre–treatment period, but there seems to be a small divergence in the trend already in 2003 and 2004 for the first treatment group. The two bottom graphs in Figure 1.2 plot the employment–to–population ratios for the treatment and control groups of men. For men, the common trend assumption seems to be satisfied as the employment rates are very stable over time or moving in the same direction for both treatment and control groups.

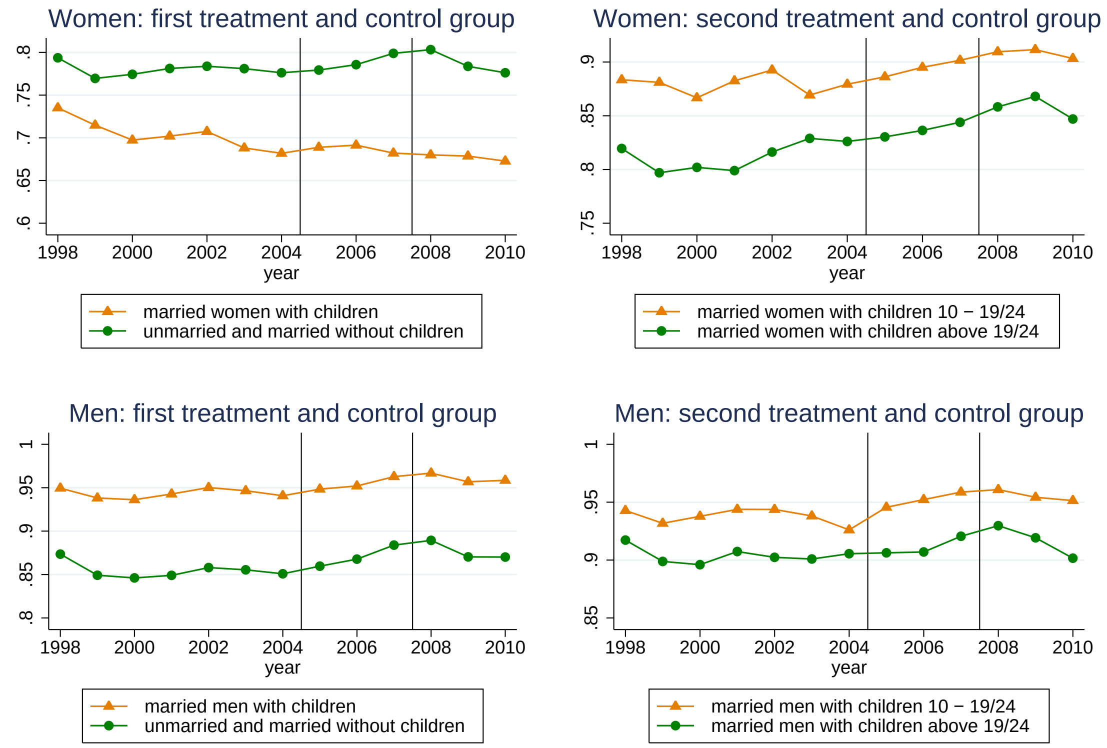
Figure 1.3 presents some evidence for the validity of the common trend assumption for the intensive labor supply measure, the average annual number of hours worked by those employed. Average hours worked have changed a little over the period for all groups of women and men analysed, with the exception of a sudden decline in 2001, which was, however, only a consequence of a change in the definition of working hours.[21]

The evidence for the validity of the common–trend assumption presented in Figures 1.2 and 1.3 is only suggestive. Since its validity seems to be more of an issue for women than for men, I provide a formal test of the common trend assumption for women as

[20] The sample is restricted to prime–aged men/women (25–54 years old).

[21] Breaks for food and rest were excluded from the working time (as a part of the unification with the EU coding), and hours worked were thus artificially decreased.

Figure 1.2: Common trend assumption: employment–to–population ratio.



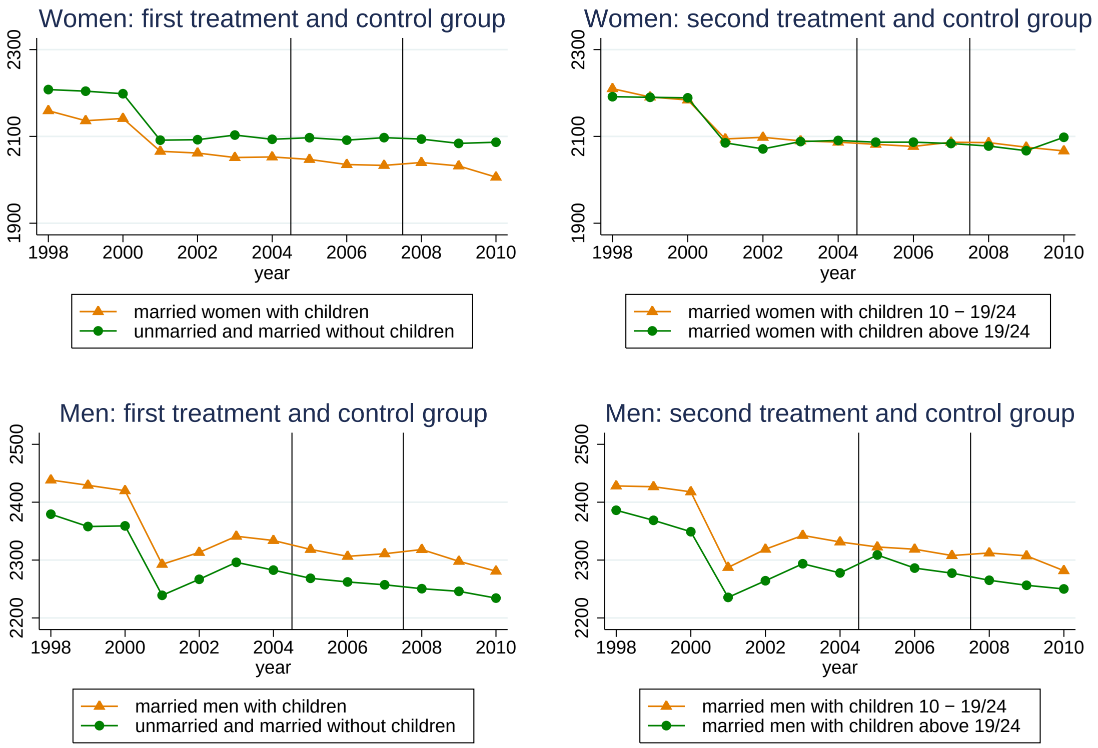
Source: EU LFS and Czech LFS, own calculations.

suggested e.g. by Heckman and Hotz (1989). Using the pre–treatment period data (1998–2004), I regress employment dummy/hours worked on a full set of year dummies, treatment dummy, interactions of treatment and year dummies, and control variables (the treatment–year interactions are reported in Appendix Table 1.9).

Although there was some uncertainty about the validity of the common trend assumption in Figure 1.2, controlling for observable characteristics helps to mitigate this problem (Appendix Table 1.9). The only treatment–year interactions that are statistically different from zero at 5% in the employment equation are those for the first treatment group and for 2000 and 2001. The validity of the first treatment and control groups might thus still be questionable, but focusing on the pre–treatment period used in the analysis (2002–2005), the common trend in employment seems to be satisfied. Results for the other three treatment and control groups show no significant departures from the common trend.

The common trend test for hours worked shows similar results as for the employment

Figure 1.3: Common trend assumption: average annual hours per worker.



Source: EU LFS and Czech LFS, own calculations.

decision. There are some significant treatment–year interactions for the first treatment and control groups, but mainly for the early pre–treatment period that is not used in the estimation. The common trend test for the other three treatment and control groups gives even better results. Therefore, the evidence on the validity of the common trend assumption is largely confirmed by the data, and even though some small departures from the common trend were found, I show in Section 1.5.1 that the analysis gives consistent results for all treatment and control groups. Moreover, I provide sensitivity analyses of the estimated results in Section 1.6, and I show that the results of these sensitivity analyses support the validity of the chosen approach.

Looking at the common trend assumption for the triple differences approach, Figure 1.5 in the Appendix plots the differences in employment and in average hours worked between the treatment and control groups in the Czech and Slovak Republics. The differences between the groups of both men and women in both countries is quite stable with small fluctuations only within a very small range of 0.05 percentage points of employment probability and 30 working hours per year.

The second identifying assumption of the difference–in–differences approach, the absence of composition changes in the treatment and control groups, could be violated if the marriage and fertility decisions of Czech couples were significantly influenced by the introduction of joint taxation or any other child–related reforms (see Section 1.3). However, empirical studies usually find a very small response on these margins (see e.g., Eissa and Hoynes 2000 or Ellwood 2000). Also, this could only compromise the validity of the first control group, because the composition of other treatment and control groups cannot be changed with fertility and marriage decisions within the given period of time.

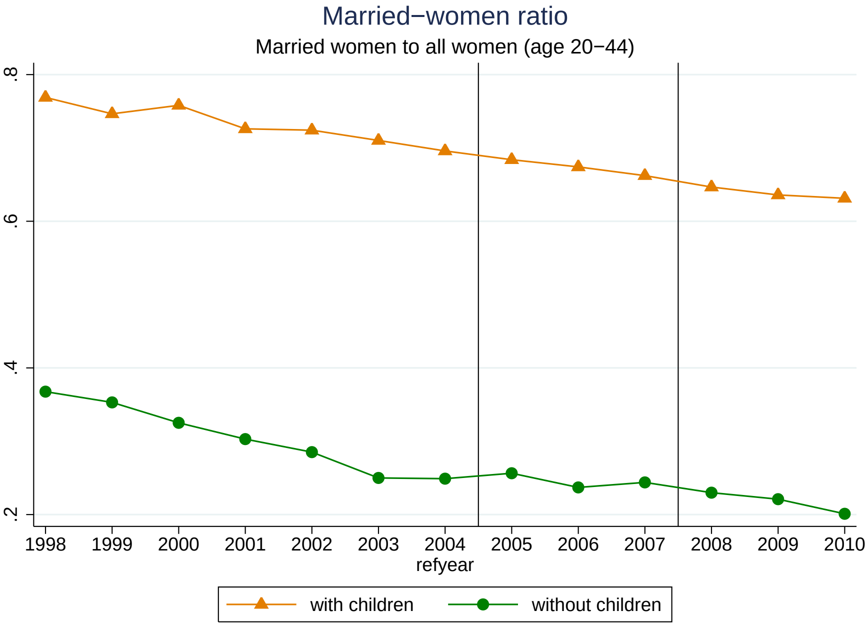
Moreover, Figure 1.4 provides evidence that the marriage decisions of Czech women were not affected by joint taxation (I focus on marital status, because it is probably easier to adjust than fertility choices). It illustrates a married–women ratio (ratio of married women to all women) for the groups of women with and without children. If there were an effect of joint taxation on marriage decisions, we would see an increase in the ratio of married women among those with children, because that would make them eligible for joint taxation. However, this is clearly not the case. On the contrary, the married–women ratio slightly increased for the group of women without children, while the trend for women with children was left unchanged.

The validity of the third and fourth treatment and control groups could also be jeopardised if children of different ages experienced different shocks to employment or education that roughly coincided with the introduction of joint taxation. This might have had an impact on the composition of the treatment and control groups and/or invalidate the common trend assumption if mothers or fathers were sufficiently responsive in their labor market decisions to the economic status of their children. Table 1.3 illustrates the share of population in education and in unemployment by age for individuals aged between 16 and 19 and between 24 and 27.[22] Clearly, the share of children in education increased quite steadily over time for all age categories, and the share of unemployed children increased slightly up until 2005, and then started decreasing for all age groups. Therefore, there were no differential shocks to education or unemployment that might compromise the validity of the treatment and control groups.

Lastly, for my estimation strategy to be valid, it is also necessary that the family taxation reform is exogenous to the outcome of interest (labor supply decisions). Among

[22] These age categories coincide with ages of children in the third and fourth treatment and control groups, to capture the possible changes in education and unemployment probabilities for these groups.

Figure 1.4: Marriage decisions by groups of women.



Source: EU Labor Force Survey, own calculations.

others, Besley and Case (2000) argue that policy actions are often purposeful in responding to economic conditions in a particular country, in which case it may be inappropriate to treat such actions as sources of exogenous variation. However, the change in family tax treatment in 2005 was implemented with the purpose of increasing tax relief for families with children, and it is very unlikely that the reform was meant to decrease the labor supply of married women or in any way affect the labor supply of married men with children.

1.4.4 Data

I use the Czech Labor Force Survey (LFS) data, which is a large sample survey covering about 60,000 Czech individuals quarterly. This dataset includes information about household structure, detailed demographic characteristics of all household members, an indicator of economic activity during the reference week (employed/unemployed/inactive), and the number of hours worked in the reference week (if employed). I use annual LFS data for three years before the introduction of joint taxation (2002–2004) and three years with joint taxation (2005–2007).

For the difference–in–difference estimation I use the original Czech LFS, but for the

Table 1.3: Share of population in education and in unemployment by age.

	2002	2003	2004	2005	2006	2007
Age	Share in education					
16	97.56%	97.68%	97.67%	97.91%	97.97%	98.09%
17	94.62%	95.79%	95.38%	96.47%	95.73%	96.96%
18	85.42%	85.60%	87.23%	86.18%	87.82%	88.37%
19	64.34%	66.12%	67.50%	69.76%	70.01%	74.41%
24	9.68%	11.48%	15.62%	17.01%	21.09%	23.01%
25	5.92%	6.84%	8.38%	9.73%	10.47%	12.52%
26	3.05%	3.98%	4.16%	5.40%	5.80%	5.80%
27	1.67%	1.60%	2.45%	2.74%	3.26%	3.18%
	Share of unemployed					
16	0.31%	0.67%	0.53%	0.74%	0.78%	0.55%
17	1.74%	1.38%	1.74%	1.33%	1.72%	1.13%
18	4.98%	5.83%	4.89%	6.35%	4.69%	2.86%
19	9.07%	9.22%	9.63%	9.09%	8.58%	5.16%
24	6.30%	7.64%	7.73%	7.08%	7.69%	3.97%
25	5.97%	6.84%	6.71%	6.45%	6.69%	4.66%
26	5.83%	5.21%	6.52%	7.19%	5.13%	4.47%
27	6.08%	7.06%	7.12%	6.70%	4.57%	3.66%

triple differences approach, where data for Slovak individuals is needed as well, I have to use the standardized EU Labor Force Survey (EU LFS). The problem is that the information available in the EU LFS is not as detailed as in the national LFS.[23] In particular, the EU LFS includes only 5–year age bands. Therefore, an accurate indicator for children up to the age of 18/26 cannot be created. Children can only be defined as those younger than 20, or full–time students younger than 25 years of age. Therefore, the treatment group in the triple difference estimation misses some eligible men/women and contains some ineligible men/women biasing the size of joint taxation effect downwards.[24]

The sample is restricted to prime–aged men/women (aged 25 to 59), who are not in full–time education. Appendix Table 2.1 reports averages of the main outcome and control variables by treatment group and treatment period calculated from the Czech

[23]EU LFS is created based on the original LFS data that are collected by national statistical offices; however, it is then processed and adjusted to correspond to the common coding scheme of the Eurostat. I use annual EU LFS data from years 2002–2007. For more information on the EU LFS, see: `http://epp.eurostat.ec.europa.eu/portal/page/portal/microdata/lfs`.

[24]Also, up until 2005, the yearly series of the EU LFS were based only on the data collected in the second quarter of the year (data for other quarters have very limited information, for example, they do not include information on marital status and the relationship between individuals within a household). Therefore, the sample for triple differences estimation is restricted to data collected in the second quarter only (to make it comparable across years).

LFS data. The employment rate of Czech married women with children is 71.5% in the period before joint taxation (2002–2004), and increases slightly to 71.9% in the period after (2005–2007). If we focus on married women with older children, the employment rate is obviously much higher (around 88%). Unmarried and childless women had a somewhat lower employment rate, and experienced a significant increase in employment probability: from 67.7% to 70.1% over the period in question. The summary statistics for men in Appendix Table 2.1 show that married men with children have a very high employment rate (around 95%), while unmarried and childless men have an employment rate of almost 10 percentage points lower. However, all groups of men experienced an increase in employment probability of a very similar magnitude (around 1 percentage point) after the introduction of joint taxation.

There is very little variation in the hours worked by employed individuals across groups of men and women; basically all groups of women worked an average number of hours per week close to 40, while men work slightly longer hours (around 43 hours per week). Most other characteristics are also pretty stable over time within each group, the main exception being education—the level of education of both men and women in the sample increased over time. Table 2.1 also confirms that while married individuals with children are quite different in some of their observable characteristics (such as level of education, number of household members, economic activity of the partner/husband) from unmarried or childless individuals, couples with children aged just below and just above the age threshold are much more similar in the observable characteristics.

1.5 Results

1.5.1 Difference–in–differences approach

In this section, I present the main estimation results of the effect of joint taxation on the extensive and intensive margins of the labor supply of married men and women with children. Table 1.4 reports the difference–in–differences coefficients (the interaction of the treatment group dummy with joint taxation years) for women, where Columns 1 to 4 correspond to the four treatment and control groups. The effect on employment decisions of married women with children is negative and significant for all treatment groups. The coefficient of -0.029 in column 1 indicates that married women with children experienced a 2.9 percentage point decline in the probability of being employed in the period of joint taxation (2005–2007), relative to unmarried women and married women without children and relative to the period before joint taxation.

Table 1.4: Difference–in–differences estimation results for women.

	(1) 1. treatment and control	(2) 2. treatment and control	(3) 3. treatment and control	(4) 4. treatment and control
Employed last week				
DID coef.	-0.029***	-0.020***	-0.028**	-0.035*
	(0.007)	(0.004)	(0.012)	(0.019)
R^2	0.271	0.160	0.115	0.244
Observations	376517	118869	16267	8772
Hours worked per week				
DID coef.	-0.368***	-0.305***	-0.551	1.359***
	(0.083)	(0.087)	(0.376)	(0.425)
R^2	0.027	0.014	0.015	0.045
Observations	262912	99628	14124	6714

Note: Standard errors (in parentheses) are clustered at the group–year level (* p<0.10, ** p<0.05, *** p<0.01). All regressions include control variables, treatment and year dummies. The first treatment and control groups compare married women with children to unmarried and childless women; the second treatment and control groups compare married women with children aged 10–17/25 to married women with children aged 18/26–30; the third treatment and control groups compare married women with children aged 16–17 vs. 18–19; the fourth treatment and control groups compare married women with children aged 24–25 vs. 26–27. Source: Czech LFS, own calculations.

The employment effect of joint taxation for the second treatment group (married women with children aged 10–17/25) is somewhat smaller than for the first treatment group (a 2 percentage point decline in the employment probability, see Column 2 of Table 1.4). Columns 3 and 4 report results of the local difference–in–differences regressions around the two age thresholds. Estimated effects at both thresholds are close to a 3 percentage points decrease in the employment probability of married women with children below the age threshold (as compared to married women with children above the age threshold).

The lower part of Table 1.4 shows the effect on hours worked. It is negative and significant at 1% for the first two treatment groups, but the effect is rather small in magnitude (the estimates suggest a decrease in hours worked per week by less than 0.4 hours). The estimated effects on hours worked for the local difference–in–differences (Columns 3 and 4) are mixed—the effect on hours worked is insignificant for married women with children below 18, and positive and significant for women with children below 26. That fact that the effect is present mainly at the extensive margin is not surprising given the low availability of jobs with flexible working hours in the Czech Republic (see e.g. Tang and Cousins 2005).

The control variables have the expected signs in all regressions (see Appendix Table 1.11): labor supply is increasing in age (but not linearly) and also in education; the presence of children of all ages decreases employment and hours worked; labor supply decreases in the number of household members; higher education of the partner leads to the higher employment probability of women, while inactivity of the partner decreases the employment probability of a woman, and non–Czech citizens are less likely to be employed, but work more hours.

The main estimation results for men are reported in Table 1.5, where Columns 1 to 4 again correspond to the four treatment and control groups. The theoretical predictions for men are ambiguous as the substitution effect of the reform motivates them to increase their labor supply and the income effect to reduce it. The results show mostly an insignificant effect on the employment probability with the exception of the fourth treatment group of married men with children aged 24–25, for whom the effect is negative and significant at 5%. This is in line with the findings of LaLumia (2008), who also found a mostly insignificant effect on labor supply of married men.

The estimated effect on hours worked by men is mostly negative and statistically significant, but the magnitude of the effect is again very small (0.1 hours per week decline for the first treatment groups, 0.4 for the second treatment group—see the lower part of Table 1.5). This is consistent with the income effect slightly outweighing the substitution effect, while the small magnitude of the effect can also be explained by the very low availability of flexible working hours in the Czech Republic.

1.5.2 Difference–in–differences by the intensity of treatment

In the Institutional background section, I illustrated that the effect of joint taxation on the female labor supply incentives varies greatly by the wife's and her husband's tax bracket. In this section, I investigate whether the effect of joint taxation on female labor supply indeed differs across groups of women based on the intensity of treatment (the size of the change in their work incentives). What matters most for the intensity of treatment is the difference between the tax bracket in which the woman's husband's income belongs

Table 1.5: Difference–in–differences estimation results for men.

	(1) 1. treatment and control	(2) 2. treatment and control	(3) 3. treatment and control	(4) 4. treatment and control
Employed last week				
DID coef.	-0.004	-0.002	0.002	-0.032**
	(0.002)	(0.004)	(0.013)	(0.013)
R^2	0.123	0.070	0.087	0.080
Observations	357948	113613	15924	7611
Hours worked per week				
DID coef.	-0.137***	-0.415***	0.324	-1.276**
	(0.019)	(0.070)	(0.297)	(0.551)
R^2	0.037	0.033	0.041	0.069
Observations	316148	104122	14656	6767

Note: Standard errors (in parentheses) are clustered at the group–year level (* p<0.10, ** p<0.05, *** p<0.01). All regressions include control variables, treatment and year dummies. The first treatment and control groups compare married men with children to unmarried and childless men; the second treatment and control groups compare married men with children aged 10–17/25 to married men with children aged 18/26–30; the third treatment and control groups compare married men with children aged 16–17 vs. 18–19; the fourth treatment and control groups compare married men with children aged 24–25 vs. 26–27. Source: Czech LFS, own calculations.

and the tax bracket in which the woman's own (actual or potential) income belongs. Unfortunately, the LFS data used in the analysis do not include any information about incomes.

To tackle this problem, I approximate the level of work income by education. Table 1.6 illustrates how the difference–in–differences coefficients for women differ by the education of a woman's husband. The estimated coefficients confirm that the effect on employment and hours worked is larger and more significant among women with more educated husbands, which is consistent with theoretical predictions. The employment effect for married women with children and a tertiary–educated husband from both the first and the second treatment group, is a 5.5 percentage points decline in employment probability (see Columns 1 and 2 in Table 1.6). The effect on hours worked is also somewhat higher for more educated husbands, but still economically insignificant (see Columns 3 and 4).[25]

Table 1.12 in the Appendix reports the difference–in–differences coefficients for women

[25] Due to small sample sizes in the third and fourth treatment and control groups, I report how the reform effect differs by the intensity of treatment only for the first and second treatment and control groups.

Table 1.6: Difference–in–differences estimation results for women by education of husband.

	(1) 1. treatment and control	(2) 2. treatment and control	(3) 1. treatment and control	(4) 2. treatment and control
	Employed last week		*Hours worked per week*	
DID coeff. interacted with:				
husband primary education	-0.010	-0.045**	0.887	0.421
	(0.012)	(0.019)	(0.606)	(0.312)
husband secondary education	-0.024***	-0.011**	-0.348***	-0.323***
	(0.007)	(0.005)	(0.085)	(0.100)
husband tertiary education	-0.055***	-0.055***	-0.653***	-0.323*
	(0.009)	(0.012)	(0.160)	(0.158)
R^2	0.272	0.161	0.027	0.014
Observations	376517	118869	262912	99628

Note: Standard errors (in parentheses) are clustered at the group–year level (* p<0.10, ** p<0.05, *** p<0.01). All regressions include control variables, treatment and year dummies. The first treatment and control groups compare married women with children to unmarried and childless women; the second treatment and control groups compare married women with children aged 10–17/25 to married women with children aged 18/26–30. Source: Czech LFS, own calculations.

interacted with the woman's and her husband's education.[26] The results generally confirm the findings from Table 1.6—the employment effect is negative and increasing with the husband's education, but also higher for more educated women. In particular, the effect is largely insignificant for primary educated women, but significant and negative for secondary and tertiary educated women. Since 81% of the primary educated women earned incomes that belonged to the first tax bracket[27] this is consistent with the theoretical predictions in Section 1.3, where the disincentive effects of joint taxation are higher for women in the second or third tax bracket than for those in the first tax bracket.[28]

[26]For some combinations of the husband's and wife's education there were too few observations to create a reliable measure of the reform effect (denoted as N/A.

[27]This percentage was calculated using the Czech SILC data for 2005.

[28]Unless the husband's income belongs to the third or fourth tax bracket. However, these combinations of tax brackets (the wife in the first and the husbands in the third or fourth tax bracket) were very uncommon for the Czech couples (the estimated share of couples with this combination of tax brackets based on the SILC data is only 5.2%).

1.6 Robustness analysis

In this section, I provide additional evidence supporting the validity of the difference–in–differences approach and the conclusions drawn from its results. In particular, I provide results of the triple differences approach with Slovak married individuals serving as additional control group and of two difference–in–differences placebo tests.

1.6.1 Triple differences approach

This section reports the estimates of the triple differences approach, which compares the treatment group of Czech married individuals with children aged 10–17/25, and two control groups—Slovak married individuals with children aged 10–17/25 and individuals with children over the age threshold (aged 18/26–30). The triple differences approach is somewhat less restricted than the difference–in–differences approach as it allows for different labor supply trends in the two countries, across the two groups of men/women, and also for differences in tastes for work between the two groups of men/women in the two countries.

The estimation results are reported in Table 1.7. The effect of joint taxation on the employment of Czech married women with children (aged 10–19/24) remains significantly negative, and thus confirms the findings of the previous section. The coefficient of -0.016 in Column 1 suggest a 1.6 percentage point decline in the employment probability of Czech married women with children aged 10–19/24 during the period of joint taxation. The magnitude of the effect is only slightly smaller compared to the difference–in–difference analysis, where the effect on this group of women was 2 percentage points.

The effect on hours worked is negative, but insignificant using the triple differences approach, so the presence of a significant effect at the intensive margin of female labor supply is not confirmed. The results for married men with children (aged 10–19/24) confirm that joint taxation had no significant effect on male employment probability. The effect on hours worked by men is also not significantly different from zero.

Further, I check if the results of the intensity–of–treatment analysis from Section 1.5.2 are supported by the triple differences approach. To do so, I interact the triple differences coefficient for women with their husband's education. The estimated effect is again highest and most significant for women with tertiary–educated husbands, which is consistent with the results in Section 1.5.2 and with the theoretical predictions in Section 1.3. The coefficient of the reform effect for women with tertiary–educated husbands is -0.043

Table 1.7: Triple differences estimation results.

	(1) Women	(2) Men	(3) Women	(4) Men
	Employed last week		*Hours worked per week*	
DIDID coef.	-0.016**	-0.008	-0.002	-0.316
	(0.006)	(0.006)	(0.125)	(0.240)
R^2	0.128	0.107	0.010	0.031
Observations	43193	38897	35914	35109

Note: Standard errors (in parentheses) are clustered at the country–group–year level (* p<0.10, ** p<0.05, *** p<0.01). All regressions include control variables, treatment, year and country dummies, and interactions of year, group, and country dummies. Source: EU LFS, own calculations.

with a standard error of 0.01, suggesting a 4.3 percentage decrease in the employment probability of treated women with highly–educated husbands.

1.6.2 Placebo tests

In this section, I repeat the difference–in–differences analysis for two hypothetical treatment and control groups of Czech married individuals. I construct these groups so that neither of them were affected by the introduction of joint taxation, but they are otherwise very similar to the groups of individuals used in the main analysis. In particular, the placebo groups are very similar to the third and fourth treatment and control groups, because they also consist of married individuals who differ only by the age of the youngest child. The first placebo test consists of a comparison of married individuals with children aged 18–20 vs. 21–23, who are not in education, while the second placebo test compares married individuals with children aged 26–28 vs. 29–31. All of these groups consist of married individuals with children above the age threshold, the analysis should thus reveal no effect of the introduction of joint taxation. Results of the two placebo tests for men and women are presented in Table 1.8.

Clearly, the effect of joint taxation on labor supply is not statistically different from zero at 5% confidence level in any of the estimated equations. This supports the claim that the results of the difference–in–differences analysis in Section 1.5.1 are indeed capturing the treatment effect of the introduction of joint taxation, and not of some other confounding factors.

Table 1.8: Results of the placebo difference–in–differences estimation.

	(1)	(2)	(3)	(4)
	Results for women		*Results for men*	
	1. placebo group	2. placebo group	1. placebo group	2. placebo group
Employed last week				
DID coef.	-0.017	-0.014*	0.007	0.010
	(0.012)	(0.007)	(0.007)	(0.020)
R^2	0.171	0.257	0.076	0.069
Observations	16342	14429	15457	11780
Hours worked per week				
DID coef.	-0.118	-0.173	-0.152	0.455*
	(0.230)	(0.429)	(0.295)	(0.218)
R^2	0.016	0.029	0.036	0.050
Observations	13198	9971	13607	10194

Note: Standard errors (in parentheses) are clustered at the group–year level (* p<0.10, ** p<0.05, *** p<0.01). All regressions include control variables, treatment and year dummies. The first placebo test compares married individuals with children aged 18–20 vs. 21–23, who are not in education; the second placebo test compares married individuals with children aged 26–28 vs. 29–31. Source: Czech LFS, own calculations.

1.7 Conclusion

The theoretical impact of joint taxation on the labor supply of secondary earners (usually women) is negative, while the impact on primary earners (usually men) is ambiguous. Despite the fact that married couples file jointly in many European countries, the magnitude of the joint taxation effect on labor supply remains unclear as the empirical literature is limited by a lack of recent policy changes with respect to family taxation.

This paper utilizes the most recent family taxation reform, the introduction of joint taxation in the Czech Republic in 2005, to investigate the labor supply effects of joint taxation on married couples. In the period 2005–2007 (inclusive) married couples raising at least one child could have used the opportunity of joint filing in the Czech Republic. I used a difference–in–difference estimation strategy with four alternative treatment and control groups to estimate the magnitude of the joint taxation effect on the labor supply of Czech married couples with children. Joint taxation in the Czech Republic was introduced as a voluntary option and the take–up rate by married couples was very high (69 to 86%). Since the take–up information is not available in the individual–level data, the estimated parameters should be interpreted as intention–to–treat effects and thus as a lower bound for the effect of mandatory joint taxation.

The results show that joint taxation indeed negatively impacts the labor supply of married women with children. Using the difference–in–differences approach, the results suggest a 2.9 percentage point decline in the employment probability of Czech married women with children compared to unmarried and childless women and compared to the period before joint taxation in the Czech Republic. The effect is a somewhat smaller (a 2 percentage points decline) among married women with older children (aged 10–17/25). The response at the intensive margin is statistically significant, but rather negligible (the results suggest a decline in the number of hours worked per week by 0.4 hours for married women with children).

To construct additional treatment and control groups, I use the age thresholds that define children in the Czech law and I apply a local difference–in–differences analysis comparing women with children just below and just above the age threshold. The results of this estimation also confirm a negative and significant effect of joint taxation on married women's labor supply at the extensive margin, of a magnitude close to 3 percentage points.

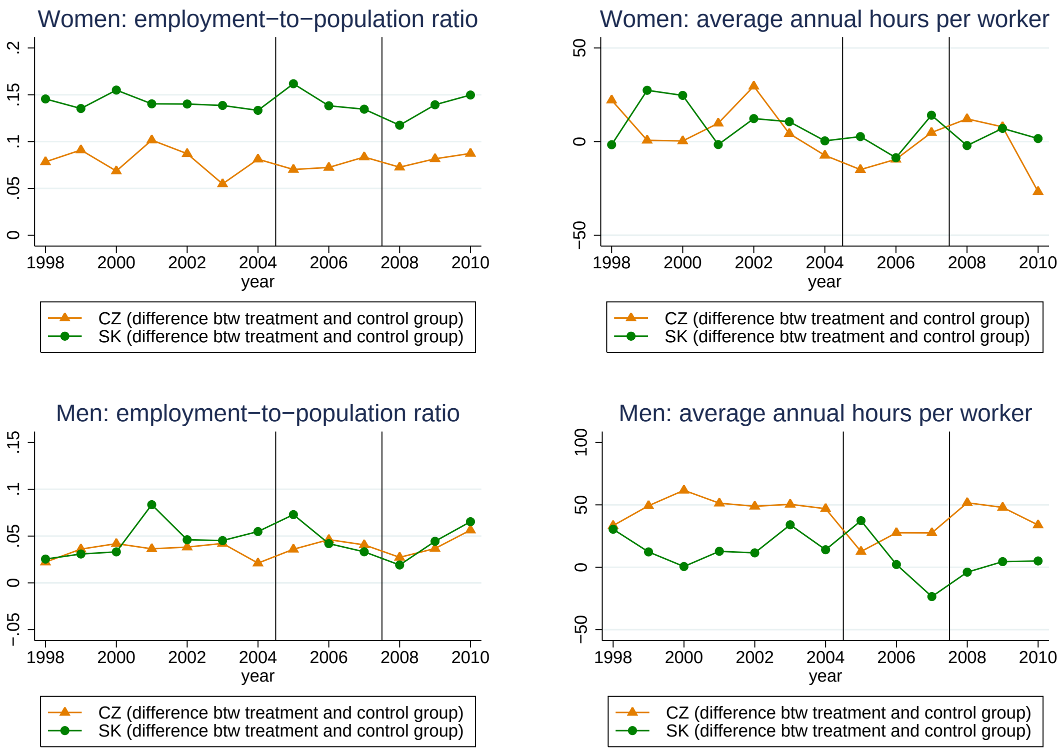
Furthermore, I take advantage of heterogeneity in the intensity of treatment caused by the joint taxation reform. The change in work incentives of married women with children varied a lot according to the difference between their husband's and their own (potential) wages. I show that those women who experienced the highest change in work incentives (women with highly–educated husbands) indeed responded with the largest decrease in employment probability, namely by 5.5 percentage points.

Results for men suggest that married men with children did not adjust their labor supply at the extensive margin in response to the introduction of joint taxation, but they decreased the hours worked slightly (by 0.1 hours per week). The insignificant effect on the employment probability is in accordance with the results of LaLumia (2008), and the small negative effect on hours worked is consistent with the income effect slightly outweighing the substitution effect.

This study shows that even in a country like the Czech Republic with high female labor force participation and relatively low elasticity of labor supply, incentives provided by the tax system have important consequences for the female labor supply. The effect is likely to be greater in other EU countries, where the labor supply of women is more sensitive to wages. Policy makers in many EU countries strive to increase female labor supply, and this paper suggests that joint taxation should not be neglected as one of its important determinants.

1.A Appendix 1

Figure 1.5: Common trend assumption: triple differences approach.



Source: EU LFS and Czech LFS, own calculations.

Table 1.9: Common trend test on pre–treatment data for women.

	(1) 1. treatment and control	(2) 2. treatment and control	(3) 3. treatment and control	(4) 4. treatment and control
Employed last week				
Treatment dummy interacted with year:				
1999	-0.002	-0.003	-0.026	-0.064*
	(0.005)	(0.007)	(0.016)	(0.035)
2000	-0.017***	-0.012*	-0.028	-0.017
	(0.005)	(0.007)	(0.017)	(0.034)
2001	-0.011**	-0.010	0.028	-0.027
	(0.005)	(0.007)	(0.019)	(0.033)
2002	-0.008	-0.008	0.034*	0.005
	(0.005)	(0.007)	(0.021)	(0.033)
2003	-0.009*	-0.010	0.043*	-0.002
	(0.005)	(0.007)	(0.022)	(0.031)
2004	-0.007	-0.010	0.013	-0.049
	(0.005)	(0.007)	(0.022)	(0.032)
R^2	0.246	0.043	0.050	0.043
Observations	297926	141421	23502	7701
Hours worked per week				
Treatment dummy interacted with year:				
1999	-0.354***	-0.256*	0.270	1.028
	(0.109)	(0.141)	(0.313)	(0.852)
2000	-0.075	-0.160	0.557	0.402
	(0.110)	(0.142)	(0.341)	(0.818)
2001	0.231**	-0.143	0.561	1.196
	(0.106)	(0.141)	(0.370)	(0.833)
2002	0.386***	0.121	0.432	-0.407
	(0.107)	(0.143)	(0.379)	(0.828)
2003	-0.016	-0.092	0.474	-1.233
	(0.109)	(0.144)	(0.412)	(0.805)
2004	0.066	-0.316**	0.861**	-0.568
	(0.107)	(0.143)	(0.402)	(0.705)
R^2	0.041	0.033	0.034	0.060
Observations	222237	122583	20398	6470

Note: Standard errors (in parentheses) are robust (* p<0.10, ** p<0.05, *** p<0.01). All regressions include treatment and year dummies and controls for region, education, and presence of children of various ages in a household. The omitted year is 1998. The first treatment and control groups compare married women with children to unmarried and childless women; the second treatment and control groups compare married women with children aged 10–17/25 to married women with children aged 18/26–30; the third treatment and control groups compare married women with children aged 16–17 vs. 18–19; the fourth treatment and control groups compare married women with children aged 24–25 vs. 26–27. Source: Czech LFS, own calculations.

Table 1.10: Summary statistics of the sample by treatment group and period.

	1. treatment and control: Married couples with children vs. unmarried and childless				2. treatment and control: Married couples with children aged 10–17/25 vs. 18/26–30			
	Control		Treatment		Control		Treatment	
	Before	After	Before	After	Before	After	Before	After
Women:								
Employed last week	0.677	0.701	0.715	0.719	0.746	0.767	0.876	0.887
Hours worked per week	40.03	40.09	39.48	39.24	39.93	40.02	40.10	39.95
Age	45.89	45.73	37.72	38.28	49.17	49.65	42.44	42.78
Primary–educated	0.193	0.165	0.082	0.067	0.212	0.192	0.098	0.076
Secondary–educated	0.723	0.741	0.798	0.796	0.742	0.759	0.785	0.797
Tertiary–educated	0.083	0.093	0.12	0.137	0.045	0.049	0.117	0.127
Men:								
Employed last week	0.838	0.85	0.942	0.951	0.872	0.884	0.932	0.946
Hours worked per week	43.47	43.21	44.73	44.38	43.48	43.58	44.64	44.44
Age	43.81	43.60	40.19	40.85	50.54	51.17	45.06	45.47
Primary–educated	0.452	0.445	0.044	0.036	0.077	0.067	0.043	0.034
Secondary–educated	0.488	0.492	0.794	0.795	0.846	0.848	0.792	0.795
Tertiary–educated	0.06	0.063	0.161	0.169	0.077	0.084	0.164	0.171
Family characteristics:								
Married	0.509	0.487	1	1	1	1	1	1
Cohabiting	0.091	0.114	0	0	0	0	0	0
No. of HH members	2.69	2.616	3.937	3.924	3.66	3.601	3.896	3.883
Children aged 0–2	0.02	0.025	0.159	0.171	0	0	0	0
Children aged 3–5	0.025	0.026	0.191	0.196	0	0	0	0
Children aged 6–9	0.044	0.043	0.295	0.279	0	0	0	0
Children aged 10–14	0.075	0.076	0.472	0.422	0	0	0.562	0.495
Children aged 15–17	0.051	0.053	0.293	0.301	0	0	0.462	0.459
Observations	113391	116191	74555	72380	21683	18344	39985	38857

	3. treatment and control: Married couples with children aged 16–17 vs. 18–19				4. treatment and control: Married couples with children aged 24–25 vs. 26–27			
	Control		Treatment		Control		Treatment	
	Before	After	Before	After	Before	After	Before	After
Women:								
Employed last week	0.79	0.805	0.884	0.882	0.71	0.766	0.877	0.881
Hours worked per week	39.83	40.36	40.12	40.13	39.94	39.77	39.71	41.05
Age	45.70	45.53	43.97	43.55	52.25	52.01	51.01	50.44
Primary–educated	0.255	0.238	0.134	0.095	0.221	0.172	0.053	0.037
Secondary–educated	0.721	0.744	0.775	0.788	0.717	0.763	0.746	0.787
Tertiary–educated	0.024	0.018	0.091	0.116	0.062	0.066	0.201	0.176
Men:								
Employed last week	0.867	0.869	0.928	0.936	0.871	0.892	0.942	0.942
Hours worked per week	43.48	42.87	44.32	44.28	43.01	43.88	43.85	43.63
Age	48.13	48.25	46.60	46.26	51.27	52.24	52.56	52.26
Primary–educated	0.117	0.09	0.047	0.042	0.057	0.053	0.028	0.009
Secondary–educated	0.83	0.883	0.814	0.804	0.838	0.827	0.683	0.691
Tertiary–educated	0.052	0.028	0.139	0.154	0.104	0.12	0.289	0.3
Partner not working	0.153	0.152	0.082	0.073	0.229	0.172	0.095	0.1
Family characteristics:								
No. of HH members	3.739	3.764	3.89	3.852	3.512	3.437	3.506	3.534
Observations	1397	1115	6829	6926	3585	3399	751	1037

Note: The table reports means of the outcome and control variables used in the regressions. The treatment period is defined as before (2002–2004), and after (2005–2007) the introduction of joint taxation. Source: Czech LFS data.

Table 1.11: Difference–in–differences estimation results for women, full specification.

	(1) 1. treatment and control	(2) 2. treatment and control	(3) 1. treatment and control	(4) 2. treatment and control
	Employed last week		*Hours worked per week*	
DID coef.	-0.029***	-0.020***	-0.368***	-0.305***
	(0.007)	(0.004)	(0.083)	(0.087)
treatment	0.016*	0.022**	-0.020	0.218**
	(0.008)	(0.007)	(0.098)	(0.096)
age	0.098***	0.128***	0.326***	0.301***
	(0.005)	(0.014)	(0.041)	(0.077)
age squared	-0.001***	-0.002***	-0.004***	-0.004***
	(0.000)	(0.000)	(0.001)	(0.001)
secondary education	0.178***	0.140***	0.846***	0.854***
	(0.005)	(0.007)	(0.081)	(0.134)
tertiary education	0.290***	0.212***	1.709***	2.075***
	(0.012)	(0.007)	(0.094)	(0.196)
children 0-2	-0.544***	.	-3.680***	.
	(0.009)	.	(0.528)	.
children 3-5	-0.234***	.	-2.045***	.
	(0.005)	.	(0.122)	.
children 6-9	-0.078***	.	-1.079***	.
	(0.004)	.	(0.089)	.
children 10-14	-0.042***	-0.014***	-0.574***	-0.544***
	(0.003)	(0.004)	(0.042)	(0.053)
children 15-17	-0.027***	-0.016***	-0.217***	-0.271***
	(0.002)	(0.003)	(0.069)	(0.066)
number of HH members	-0.011***	-0.030***	-0.009	-0.068
	(0.003)	(0.002)	(0.045)	(0.047)
married	0.012	.	-0.064	.
	(0.008)	.	(0.138)	.
cohabiting	-0.012*	.	0.135	.
	(0.006)	.	(0.254)	.
secondary education	0.050***	0.085***	0.011	-0.304
of partner	(0.003)	(0.007)	(0.153)	(0.221)
tertiary education	0.059***	0.106***	0.045	-0.198
of partner	(0.005)	(0.009)	(0.201)	(0.200)
inactive partner	-0.135***	-0.115***	-0.006	0.011
	(0.005)	(0.007)	(0.144)	(0.167)
EU nationality	-0.021	-0.137***	0.489*	0.703
	(0.023)	(0.042)	(0.267)	(0.550)
non-EU nationality	-0.078***	-0.195***	2.206***	-0.303
	(0.017)	(0.038)	(0.530)	(0.761)
R^2	0.271	0.160	0.027	0.014
Observations	376517	118869	262912	99628

Note: Standard errors (in parentheses) are clustered at the group–year level (* p<0.10, ** p<0.05, *** p<0.01). All regressions include regional and year dummies. The first treatment and control groups compare married women with children to unmarried and childless women; the second treatment and control groups compare married women with children aged 10–17/25 to married women with children aged 18/26–30. Source: Czech LFS, own calculations.

Table 1.12: Difference–in–differences estimation results for women by couple's education.

	(1) 1. treatment and control	(2) 2. treatment and control	(3) 1. treatment and control	(4) 2. treatment and control
	Employed last week		*Hours worked per week*	
DID coef. for couple's education:				
wife primary	0.007	-0.100**	2.570***	1.248**
husband primary	(0.013)	(0.036)	(0.768)	(0.567)
wife primary	0.030	-0.012	-0.355**	-0.515**
husband secondary	(0.017)	(0.012)	(0.135)	(0.233)
wife primary	N/A	N/A	N/A	N/A
husband tertiary	N/A	N/A	N/A	N/A
wife secondary,	-0.018	0.019	-0.089	-0.053
husband primary	(0.014)	(0.012)	(0.604)	(0.700)
wife secondary	-0.025***	-0.011**	-0.338***	-0.295**
husband secondary	(0.006)	(0.005)	(0.086)	(0.111)
wife secondary	-0.035***	-0.042***	-0.273	-0.294
husband tertiary	(0.009)	(0.012)	(0.204)	(0.253)
wife tertiary	N/A	N/A	N/A	N/A
husband primary	N/A	N/A	N/A	N/A
wife tertiary	-0.057***	-0.009	-0.453**	-0.521*
husband secondary	(0.012)	(0.006)	(0.182)	(0.259)
wife tertiary	-0.077***	-0.070***	-0.997***	-0.312*
husband tertiary	(0.013)	(0.014)	(0.128)	(0.170)
R^2	0.272	0.161	0.027	0.015
Observations	376517	118869	262912	99628

Note: Standard errors (in parentheses) are clustered at the group–year level (* p<0.10, ** p<0.05, *** p<0.01). All regressions include control variables, treatment and year dummies. For some combinations of the couple's education there were too few observations to create a reliable measure of the reform effect (denoted as N/A). The first treatment and control groups compare married women with children to unmarried and childless women; the second treatment and control groups compare married women with children aged 10–17/25 to married women with children aged 18/26–30. Source: Czech LFS, own calculations.

Chapter 2

Tax and Transfer Policies and the Female Labor Supply in the EU

2.1 Introduction

The impact of tax–benefit systems on female labor supply has important consequences for optimal design of tax and transfer policies. Labor supply elasticities have been widely studied in the economic literature, with a major challenge of this literature being the endogeneity of income.[1] The literature on the responsiveness of labor supply decisions to tax and benefit changes can be separated into three main groups: structural models, reduced–form estimation, and grouped data estimation. Most labor supply elasticity estimates come from the structural literature that builds on a family labor supply model (see e.g. van Soest 1995, Hoynes 1996, Blundell et al. 2000, Bargain, Orsini, and Peichl 2014). A second group of studies uses a specific tax or transfer reform in the reduced–form estimation of labor supply responsiveness (see e.g. Eissa and Liebman 1996, Meyer and Rosenbaum 2001, Saez, Matsaganis, and Tsakloglou 2012). Finally, the grouped data literature identifies labor supply elasticities by estimating group–average regressions over a long time period (see e.g. Blundell, Duncan, and Meghir 1998, Devereux 2004, Blau and Kahn 2007, Causa 2010).

Researchers usually seek exogenous variation in income provided by tax and transfer

[1] Both labor and non–labor income are potentially endogenous to the labor supply. People have different unobserved characteristics (taste in leisure activities, ability, willingness to work hard, etc.) that affect their probability of being employed, their wages, and the size of non–labor income.

reforms or by non–linearities in tax–transfer schedules. However, modeling tax–benefit systems for more than one country in a harmonized way has been largely limited by the complex nature of tax and transfer schedules. Therefore, the literature is highly concentrated on the US, the UK, and developed economies of Western Europe, while there is little evidence for other countries, including the new EU member states. Moreover, the estimated magnitudes of female labor supply elasticities vary greatly across studies (for a survey, see Blundell and Macurdy 1999, Keane 2011, Meghir and Phillips 2008).

This study offers several contributions to the literature on female labor supply elasticity. Using the tax–benefit microsimulation model EUROMOD, I estimate the effect of tax–benefit policies on female labor supply, based on a sample of 26 European countries in 2005–2010. This allows me to study the responsiveness of female labor supply across countries and groups of women in a comparable way, and at the same time to control for both time–invariant and time–varying country–level unobserved factors. Further, this paper uses a new approach to deal with the endogeneity of income by using a group–level simulated instrumental variable based on a fixed EU–wide sample of women.

This study focuses on the extensive margin of female labor supply, because the responsiveness of female labor supply has been found to be driven mainly by participation choices (Blundell, Duncan, and Meghir 1998, Keane 2011).[2] Unlike previous literature, which uses net wage and non–labor income as the main explanatory variables, the main explanatory variable in this study is a measure of the extensive margin work incentives—the participation tax rate (PTR). The PTR is defined as a proportion of lost earnings that is compensated for by lower taxes and higher benefits when not in paid work, and it thus describes the (dis)incentives provided by the tax–transfer system for the participation decision. The use of the PTR allows me to capture the joint effect of taxes and transfers on female labor participation decisions and to deal with the endogeneity and measurement error in income by using a simulated instrumental variable.[3]

The instrumental variable for the PTR used in this study exploits variation in the PTR driven by changes in tax–transfer policies, while it eliminates the variation in the PTR caused by behavioral responses to these tax–transfer changes. In particular, the individual–level PTR is instrumented with a group–level measure of tax and transfer

[2]In accordance with previous literature, I refer to the decision to work or not as the participation decision (see, e.g. Meghir and Phillips 2008). It should not be confused with labor force participation (being in a labor force or not, including the unemployed).

[3]Note that the use of the participation tax rate in the labor supply equation itself does not solve the problem of income endogeneity, because the PTR is a function of family income.

systems based on a fixed EU–wide sample of women. This sample was created from a pooled dataset of all EU countries by taking a random sample of approximately 27,000 women. The instrumental variable (IV) for a woman with given characteristics,[4] living in country c and year t is calculated as an average PTR of women from the fixed EU–wide sample who have the same characteristics and whose PTR is computed based on the tax–transfer system of country c in year t. Therefore, the only variation in the IV stems from differences in tax and transfer policies across EU countries, over time, and across groups of women. This instrumental variable approach builds on the simulated IV approach used in the health economics literature (Currie and Gruber 1996, Cutler and Gruber 1996), and is also related to the simulated IV of Moffitt and Wilhelm (2000), Gruber and Saez (2002), and Dahl and Lochner (2012)[5] and to the grouped data literature.

This paper takes advantage of recent developments in multinational microsimulation models that allow researchers to model the tax and benefit systems of a large set of countries in a comparable way. It uses the EU–wide microsimulation model EUROMOD[6] to calculate participation tax rates at the individual level for 26 EU countries in 2005–2010. The rich structure of the data enables to study the heterogeneity in female labor supply responsiveness across countries and groups of women while controlling for time–invariant country–specific characteristics (such as culture and informal institutions), and also for time–varying country–level unobserved factors (such as country–level economic shocks, changes in preferences for work and family policies).

Multinational microsimulation models have so far been used mainly to describe the differences in the tax and transfer systems across countries, and to my knowledge, Bargain, Orsini, and Peichl (2014) is the only study that uses a multinational microsimulation model in the estimation of labor supply elasticity, and is thus closest to the present study. They use the microsimulation models TAXSIM and EUROMOD to compare labor supply elasticities of men and women in the U.S. and 17 European countries. Compared to Bargain, Orsini, and Peichl (2014), this study takes advantage of a newer version of the

[4]Characteristics include education level, the presence of children of different ages, and marital status.

[5]These studies simulate instrumental variables based on the pre–reform characteristics of the affected individuals to which they apply the post–reform tax and transfer schedules. My instrumental variable works in a similar way, but I apply the tax and transfer schedules to the fixed sample of individuals with similar characteristics in order to minimize the effect of composition changes across countries and time. See Section 2.2.2 for details.

[6]EUROMOD is a tax–benefit microsimulation model for all EU member states. In this paper, EUROMOD version F6.0+ is utilized. EUROMOD is maintained, developed and managed by the Institute for Social and Economic Research (ISER) at the University of Essex, in collaboration with national teams from the EU member states. See https://www.iser.essex.ac.uk/euromod.

EUROMOD model, which includes a larger sample of countries (mainly from the post–communist countries) and a much longer time span, and I also take a different estimation approach. My methodology is based on a reduced–form estimation combined with an instrumental variable approach, while Bargain, Orsini, and Peichl (2014) use a structural model. An advantage of the present study over the structural models is that it does not require any assumptions on preferences (including the form of utility function and the choice set for working hours). However, it does require a sufficient amount of changes in the tax and transfer policies for identification. In Section 2.2.2, I argue that there were sufficient policy variation in the EU countries between 2005 and 2010 for the approach to be effective.

My results suggest that a 10 percentage point increase in the participation tax rate decreases the probability of female employment by 2 percentage points. The implied participation elasticity with respect to the PTR is 0.08—a 10% increase in the participation tax rate decreases the female employment rate by 0.8%. I also analyze the heterogeneity of the response across groups of women with different characteristics and find that the effect is substantially higher for women with secondary education (an elasticity of 0.16, compared to an elasticity of 0.04 for primary and tertiary–educated women) and women with small children. Consistent with previous findings, the highest elasticity is found for single mothers (0.32). Finally, the use of a multinational microsimulation model allows for a comparison of female participation elasticities across groups of countries in a harmonized way. I find that the responsiveness does differ substantially across countries, and that the countries with the lowest female employment rates have the highest participation elasticity.

2.2 Methodology

2.2.1 The participation tax rate

This section introduces basic notations and explains the role of the participation tax rate in female participation decisions. Let us assume that each woman has a fixed earnings potential e_w^p and fixed costs of work q_w (including a disutility from work, a value of lost home production, child care costs, etc.). She chooses between working ($e_w = e_w^p$) and not working ($e_w = 0$) to maximize the household's utility.[7]

The effect of taxes and transfers on family income is captured by a tax–transfer function $T(e_m, e_w, \rho)$, which represents net taxes paid as a function of both spouses' earnings (e_m denotes earnings of the woman's spouse) and parameters of the tax–transfer system (ρ). Therefore, the female participation decision is based on a comparison of costs of work and net gain from entering the labor market, which is defined as gross earnings less net taxes that the woman has to pay while doing paid work on top of net taxes that she pays out of work. Therefore, the woman decides to enter the labor market if:

$$q_w \leq e_w^p - [T(e_m, e_w^p, \rho) - T(e_m, 0, \rho)]. \tag{2.1}$$

The participation decision can then be expressed in terms of the participation tax rate:

$$PTR \equiv \frac{[T(e_m, e_w^p, \rho) - T(e_m, 0, \rho)]}{e_w^p} \leq \frac{e_w^p - q_w}{e_w^p}, \tag{2.2}$$

where the PTR describes the proportion of lost earnings that is compensated by lower taxes and higher benefits when not in paid work.

2.2.2 Estimation approach

The model from the previous section provides a basis for the estimation approach used in this paper. I estimate the effect of a widely used work incentive measure—the participation tax rate—on the labor supply decisions of women. The participation equation has the following form:

$$Empl_{ict} = \alpha PTR_{ict} + \beta' X_{ict} + \gamma_t + \gamma_c + (\gamma_{ct}) + \epsilon_{ict}, \tag{2.3}$$

where $Empl_{ict}$ is the employment dummy, PTR_{ict} is the participation tax rate, and X_{ict} represents the set of observable characteristics including age, education, marital status, number of household members, dummy variables for the presence of a spouse, for children of certain ages (children aged 1, 2, 3, 4–5, 6–9, 10–15, and no children below 16),

[7] Married and cohabiting women are assumed to be secondary earners; their labor force participation decisions follow their spouses' decisions, while single women are the primary and only potential earner in a household. I also assume here that spouses pool their resources, which is a standard assumption of a unitary model. Though some studies question this assumption, the unitary model is still widely used in the labor supply literature (see e.g. Blundell, Pistaferri, and Saporta-Eksten 2012) and some recent empirical studies have supported the validity of a unitary model (see, e.g. Bargain, Orsini, and Peichl 2014).

and for elderly household members, and characteristics of spouse if present (education and economic status). I also include country fixed effects (γ_c) and year fixed effects (γ_t), while in most specifications all country–year fixed effects (γ_{ct}) are included. Therefore, I allow for changes in the country–specific fixed effects, which capture unobserved country–specific tastes for work, cultural norms, gender–role attitudes, labor market conditions, and family policies.

I deal with possible endogeneity and measurement error in the PTR by using a simulated instrumental variable.[8] The instrument for the PTR represents a group–level measure of the tax–transfer work incentives which is created based on a fixed sample of women from the whole EU. This method builds on the simulated instrument approach used in the health economics literature (Currie and Gruber 1996, Cutler and Gruber 1996), and is also related to the simulated IV used in the literature on responsiveness towards tax and transfer changes (Moffitt and Wilhelm 2000, Gruber and Saez 2002, Dahl and Lochner 2012).

The instrumental variable for the PTR is created in three steps. First, I take a random sample of 27,000 women (denoted by a subscript j) from the pooled sample of the 26 EU countries in 2007.[9] The first step provides a sample of women with fixed demographic characteristics and fixed income distribution. Second, I calculate the participation tax rate PTR_{jct} for each woman j from this fixed EU–wide sample applying country c and the year t's tax and transfer system. I repeat this PTR calculation for each country–year cell. Therefore, for each woman in the fixed EU–wide sample, I have 126 calculated PTRs, where each PTR corresponds to one country–year cell.[10] To avoid problems with income–level differences across the EU countries which might negatively affect the calculated PTR_{jct}, I adjust incomes of women from the fixed EU–wide sample to correspond to the level of incomes in country c.[11]

[8] The participation tax rate is a function of a woman's and her husband's income (see equation 2.2). Therefore, it can be affected by the standard endogeneity and measurement error problems of income in the labor supply equation.

[9] A sample of 27,000 women seems sufficiently large to provide a reasonably strong IV and EUROMOD is not meant to work with much larger samples. I have also conducted a robustness check with a substantially smaller sample of 17,000 women, but the results were largely unchanged.

[10] There are 126 country–year cells used in the estimation, because not all 26 countries are observed for all 6 years between 2005 and 2010. For details, see Section 2.3.1.

[11] I assign each woman in the fixed EU–wide sample a quantile in the income distribution of her own country, and then change her income to correspond to the average income in that quantile, but in the income distribution of country c. I create very detailed income distributions with 400 income quantiles in each country. I also adjust incomes of all household members the same way, because their incomes potentially affect the PTR computation as well.

Third, the instrumental variable for a woman i from group g, country c, and year t is constructed as an average PTR_{jct} of women from the fixed EU–wide sample who belong to group g.[12] Therefore, the only variation in this group–level IV stems from variation in tax and transfer systems across EU countries, over time, and across groups of women. The possible endogeneity and measurement error in the PTR are filtered out using the sample of women with fixed characteristics and fixed income distribution.

The simulated instrument is used in a 2SLS estimation described by the following equations:

$$PTR_{ict} = \lambda PTR\_IV_{ict} + \theta' X_{ict} + \gamma_t + \gamma_c + (\gamma_{ct}) + u_{ict}, \tag{2.4}$$

$$Empl_{ict} = \delta \widehat{PTR}_{ict} + \phi' X_{ict} + \gamma_t + \gamma_c + (\gamma_{ct}) + e_{ict}, \tag{2.5}$$

where PTR_IV_{ict} is the instrumental variable for the PTR, $\widehat{PTR}_{ict}$ denotes the predicted PTR from the first stage regression, and δ denotes the coefficient of interest.

The identification of female labor supply elasticities in this paper is based on tax and transfer changes that took place in the 26 EU countries between 2005 and 2010. Indeed, there were several reforms of tax and transfer schedules in this time period including some major reforms of tax systems—tax base allowances reform in Belgium in 2008, a flat tax reform in the Czech Republic in 2008, tax system changes in Denmark in 2010, an increase in the number of tax brackets in Spain in 2007, etc. Benefit schedules in the EU countries also underwent several important changes including the introduction of an allowance for school children in Belgium in 2006, reforms of housing and child benefits in the Czech Republic in 2007 and 2008, increased generosity of the universal child benefit and the reform of education benefits in Germany in 2008 and 2009, extensions to the large family benefit in Greece in 2006 and 2008, the introduction of a Solidarity labor income benefit in France in 2009, the reform of child benefit in Lithuania in 2009, etc.

[12] There are in total 30 groups that are defined based on three educational categories (primary, secondary, and tertiary education), five categories according to the presence of children of various ages (children aged 1–3, 4–5, 6–9, 10–15, and no children below 16) and two categories by marital status (married and unmarried). Therefore, the IV for a married childless woman with tertiary education living in Germany in 2008 is calculated as an average PTR of women from the fixed EU–wide sample who are also married, childless, and tertiary educated, and whose PTR is calculated based on the German tax–transfer system in 2008.

2.3 Data and microsimulation of taxes and benefits

2.3.1 Data and sample selection

The empirical analysis makes use of the tax–benefit microsimulation model EUROMOD, version F6.0+. The EUROMOD model is largely based upon harmonized EU–SILC data[13] (that are further adjusted for microsimulation purposes) combined with a detailed tax–benefit calculator. The model utilizes detailed information on household composition, characteristics of household members, and their incomes from the micro data, and creates common definitions of income concepts and assessment units to allow for a very detailed and harmonized micro–level calculation of taxes and benefits (for details on the EUROMOD project, see Sutherland and Figari 2013). This makes EUROMOD a very suitable instrument for computing participation tax rates in a harmonized way for the EU countries.

The EUROMOD model covers all 27 countries of the EU, but I exclude Malta from the analysis, because the Maltese data have serious shortcomings.[14] I utilize tax–transfer schedules that were in force from 2005 to 2010 and are available in the EUROMOD, version F6.0+. The EUROMOD model covers some countries only for 2006–2010 and some countries only for 2007–2010. Moreover, while EUROMOD computes taxes and transfers for all the above years, the EUROMOD input data are available only for selected years. Computation of taxes and transfers for years that do not have the corresponding input data is based on data from previous years with updated incomes. An overview of country–year cells, for which the tax–transfer computations are available and for which the EUROMOD input data are available, is provided in Appendix Table 2.5.[15]

These country–year combinations, which have tax–transfer computations in EUROMOD but do not have available input data, cannot be directly used in the estimation, because actual participation decisions of women for these country–year cells are not observed. However, the EUROMOD can be used to calculate the participation tax rates

[13] For most countries EU–SILC UDB data are used for microsimulation, but for some countries national SILC data are utilized, while the Family Resource Survey data are used for the UK.

[14] Maltese data does not include exact age information, but report age only in 5–year age bands, which is a serious limitation for female labor supply analysis, mainly because we cannot identify the exact age of children in a family.

[15] Official country abbreviations are used throughout the paper: Austria (AT), Belgium (BE), Bulgaria (BG), Cyprus (CY), Czech Republic (CZ), Germany (DE), Denmark (DK), Estonia (EE), Spain (ES), Finland (FI), France (FR), Greece (GR), Hungary (HU), Ireland (IE), Italy (IT), Lithuania (LT), Luxembourg (LU), Latvia (LV), Netherlands (NL), Poland (PL), Portugal (PT), Romania (RO), Sweden (SE), Slovenia (SI), Slovakia (SK), United Kingdom (UK).

for all available country–year combinations, and then these participation tax rates can be assigned to individuals in the EU–SILC data, where the participation decisions are available. Participation tax rates computed within the EUROMOD are imputed to the EU–SILC data based on reported incomes and household characteristics using propensity score matching.[16] The imputation should be very precise given the fact that the PTR is merely a function of incomes and other observable characteristics of individuals in a household.

Nevertheless, the quality of matching is examined in Section 2.4.3, where estimation results based on the EUROMOD data are compared to those based on the EU–SILC data with imputed participation tax rates. Since the quality of matching is indeed good, the main results presented in the paper are those based on the EU–SILC data with the imputed PTR. This allows me to take advantage of all available country–year cells in EUROMOD and substantially improves the identification strategy.

I restrict the sample to prime–aged (aged 25–55) women, and I exclude women in full–time education, pensioners, disabled, women with a new–born child (younger than 1 year of age),[17] and those with missing values for education. I also exclude the self–employed from the analysis (all women who have more than 30% of their work income from self-employment), because the quality of reporting of self–employment income in the micro–data sources is generally limited and varies widely across countries (Immervoll 2004). Excluding self–employed women is a common practice in the majority of the female labor supply elasticities papers (see, e.g. Bargain, Orsini, and Peichl 2014). The analysis includes both women living in couples (married or cohabiting) and single women.

2.3.2 Participation tax rate calculations

The participation tax rate is defined as the difference between net taxes paid when the woman works and when she does not work over her gross wage, while the economic status and incomes of all other household members are fixed. Therefore, to calculate the PTR,

[16] The propensity score matching procedure matches women in the EUROMOD data with those in EU–SILC within each country–year cell based on their income, marital status, income of the partner (if present), dummy variables for presence of children of various ages (children aged 1, 2, 3, 4–5, 6–9, 10–15, and no children below 16) and a dummy for elderly household members. Each woman in the SILC data is assigned a closest neighbor from the EUROMOD data and the corresponding PTR is imputed. To maintain consistency, the PTR is imputed to the EU–SILC data even for those years for which the EUROMOD input data are available.

[17] Children aged 0 are dropped from the EUROMOD dataset in order to align demographic variables with the income reference period for the computation of benefits (in most countries, the income reference period of the data is the calendar year preceding the survey).

I need to compute taxes and benefits for all household members for two hypothetical scenarios—when the woman works and when she does not work. For the non–working women, this requires some assumptions about their potential earnings. I impute monthly wages for all women (both working and non–working) using Heckman's two step procedure.[18]

EUROMOD is then used to calculate monthly income taxes, social security and health contributions paid, and welfare benefits received for all household members for the situation of the woman working (based on predicted wage) and not working (zero wage). Benefits included in the PTR computations consist mainly of social assistance benefits (targeted to very low income households), child–related benefits, and housing benefits.[19] Computed taxes, contributions, and welfare benefits and imputed monthly wages are then used in the PTR calculation (see equation 2.2). The same procedure is applied to calculate the PTR for the fixed EU–wide sample of women.

Sample summary statistics of the employment rate, the PTR and the IV for the PTR by country are reported in Table 2.1. There are in total over 433,000 women from 26 countries in the sample. The average employment rate of prime–aged women in the sample is 82.5%, but there are large differences across countries, with Scandinavian countries having an employment rate over 90% and Southern Europe with relatively low employment rates (close to 60% or 70%). The average participation tax rate in the sample is 30.2%, but again the PTR differs greatly across countries, with Belgium, Denmark, and Slovenia having the highest average participation tax rates (over 40%), and Cyprus, Greece, and Spain having the lowest average PTR, not exceeding 20%. The within–country variations in the PTR are also substantial and are mainly caused by the presence of means-tested benefits and progressive income tax.[20]

[18]The wage regression adjusted for selection term is run for each country and year separately to allow for different determinants of wages across countries and over time. The selection term is identified using dummies for the presence of children of different ages in the household. Other explanatory variables in the monthly wage regression include education, age, marital status, and nationality.

[19]Public pension benefits are ignored in the present study, because its focus is on prime–aged women. I also exclude maternity and parental leave benefits and unemployment benefits from the PTR computation, because the EUROMOD model includes these benefits only in a few countries (eligibility for these benefits often depends on employment history, which is not available in the data). Moreover, unemployment benefits represent only temporary income replacement, and I am more interested in medium to long term work incentive effects.

[20]For example, there are means-tested child benefits, education benefits, and social assistance benefits in Germany, which in combination with a progressive income tax system that treats married couples jointly (thereby increasing the marginal tax rates of secondary earners) creates a system with quite high and much dispersed participation tax rates for women. Lithuania provides a good example of a country with a participation tax rate that has low variance. Lithuania applies a flat tax rate system to personal income, and the only means–tested benefit is the social assistance benefit for very low income households.

Table 2.1: Summary statistics of employment rate and the PTR by country

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	Employment rate		PTR		IV for PTR		Observations
	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.	
AT	0.838	0.368	0.367	0.132	0.404	0.080	9,840
BE	0.835	0.371	0.423	0.105	0.404	0.063	14,689
BG	0.901	0.298	0.242	0.135	0.236	0.065	8,781
CY	0.775	0.418	0.109	0.152	0.106	0.071	5,385
CZ	0.837	0.370	0.308	0.116	0.350	0.053	22,042
DE	0.844	0.363	0.389	0.118	0.368	0.059	20,234
DK	0.981	0.136	0.468	0.136	0.550	0.066	9,550
EE	0.895	0.306	0.234	0.106	0.245	0.030	13,917
ES	0.732	0.443	0.175	0.091	0.165	0.054	34,864
FI	0.954	0.210	0.309	0.074	0.260	0.040	14,613
FR	0.905	0.294	0.330	0.126	0.347	0.088	21,315
GR	0.625	0.484	0.115	0.098	0.050	0.032	13,364
HU	0.825	0.380	0.326	0.109	0.365	0.094	20,729
IE	0.712	0.453	0.235	0.160	0.260	0.041	5,969
IT	0.684	0.465	0.249	0.087	0.165	0.048	46,606
LT	0.894	0.308	0.250	0.084	0.242	0.042	11,780
LU	0.740	0.439	0.394	0.183	0.427	0.063	9,585
LV	0.904	0.294	0.298	0.050	0.299	0.027	11,185
NL	0.886	0.318	0.316	0.102	0.367	0.049	22,220
PL	0.783	0.412	0.311	0.090	0.309	0.034	26,860
PT	0.818	0.386	0.270	0.139	0.375	0.091	9,373
RO	0.706	0.456	0.313	0.097	0.304	0.027	11,245
SE	0.963	0.189	0.334	0.106	0.324	0.036	13,848
SI	0.953	0.213	0.418	0.060	0.484	0.062	25,195
SK	0.905	0.293	0.327	0.181	0.347	0.067	14,643
UK	0.825	0.380	0.348	0.176	0.230	0.049	15,775
Total	0.825	0.380	0.302	0.138	0.299	0.123	433,607

Notes: The sample includes women aged 25–55, who are not in full–time education, are not pensioners, disabled, or self–employed, and do not have a child younger than 1. The number of observations for each country differs due to differences in sample sizes, and also due to the different number of years covered in different countries (see Appendix Table 2.5).
Source: EUROMOD and EU–SILC data (2005–2010), own calculations.

Summary statistics of the instrumental variable for the PTR are reported in columns 5 and 6 of Table 2.1. The mean of the IV follows quite closely the mean of the PTR in each country, which confirms that the IV captures most of the cross–country variation in the participation tax rates. The instrumental variable for the PTR has substantially smaller standard deviations than the PTR, because the IV varies only at group–level (while the PTR has an individual–level variation).

Finally, evidence of time variation in the PTR is provided in Appendix Figure 2.1. It illustrates changes in the distribution of the PTR over time separately for each country by plotting a box plot of the PTR for each country–year cell. Clearly, the distribution of the PTR changed substantially over time in all countries, with the largest changes coinciding with some of the major tax or transfer reforms—a tax base allowances reform in Belgium in 2008, a flat tax reform in the Czech Republic in 2008, tax system changes in Denmark in 2010, the introduction of Solidarity labor income benefit in France in 2009, an increase in the number of tax brackets in Spain in 2007, etc.

2.4 Results

2.4.1 Female labor supply responsiveness to tax–transfer changes

First, I report results of the first stage regressions, which indeed confirms the strength of the instrumental variable (see Table 2.2). The instrumental variable is highly significant in both the specification with and without country–year fixed effects. R^2 exceeds 0.4 and the first stage F statistic is very high (above 700), so that the null hypothesis of a weak instrument is rejected (Stock, Wright, and Yogo 2002).

Table 2.2: First stage regression results

	(1)	(2)
	Dependent var.: PTR	
PTR_IV	0.601***	0.602***
	(0.013)	(0.014)
country–year fixed effects	no	yes
R^2	0.400	0.406
F	1107.431	710.403
Observations	433,607	433,607

Notes: All regressions include a full set of country dummies, year dummies, and control variables. Standard errors in parentheses are clustered at country–year–group level, where groups are defined by education, presence of children of various ages, and marital status (* p<0.10, ** p<0.05, *** p<0.01). Source: EUROMOD model and EU-SILC data (2005–2010), own calculations.

The main results of the responsiveness of female participation towards the PTR are reported in Table 2.3. The OLS effect of the participation tax rate on an employment decision is reported in the first two columns of Table 2.3. The effect is negative and significant at 1% and suggests that an increase of the PTR by 10 percentage points decreases employment probability by 0.6 percentage points. The implied OLS elasticity of participation decision towards the PTR is 0.02.[21]

Table 2.3: Estimates of the participation equation

	(1)	(2)	(3)	(4)
	Dependent var.: employment dummy			
	OLS		2SLS	
Panel A. Estimation results				
PTR	-0.060***	-0.061***	-0.188***	-0.208***
	(0.009)	(0.008)	(0.039)	(0.030)
country–year fixed effects	no	yes	no	yes
R^2	0.174	0.175	0.173	0.173
Observations	433,607	433,607	433,607	433,607
Panel B. Elasticity of employment to PTR				
Implied elasticity	-0.022	-0.022	-0.069	-0.076

Notes: All regressions include a full set of country dummies, year dummies, and control variables. Standard errors in parentheses are clustered at country–year–group level, where groups are defined by education, presence of children of various ages, and marital status (* p<0.10, ** p<0.05, *** p<0.01). Panel A reports estimation results of equations 2.3 (OLS) and 2.5 (2SLS). Panel B reports the corresponding elasticity of employment to the PTR at the mean of independent variables.
Source: EUROMOD model and EU-SILC data (2005–2010), own calculations.

The 2SLS approach also implies a significant negative impact of the PTR on employment, but the magnitude of the estimated PTR effect is more than three times higher than in the OLS estimation, which confirms the presence of attenuation bias caused by measurement error and endogeneity of income in the OLS estimation.[22] The 2SLS estimation implies that an increase in the PTR by 10 percentage points decreases employment prob-

[21] All reported elasticities are elasticities at the mean of the independent variables. They are defined as the corresponding PTR coefficient multiplied by the mean of the PTR over the mean of an employment rate.

[22] The income endogeneity in the OLS estimation is also likely to cause a downwards bias. In most countries, the participation tax rate increases with the woman's and her husband's wages, which are both positively correlated to the woman's employment probability (through assortative matching). Therefore, the OLS estimates are likely to be biased downwards, because they include this endogenous positive correlation between the woman's PTR and her employment probability.

ability by 2 percentage points (see columns 3 and 4 of Table 2.3). The implied elasticity of labor supply with respect to the participation tax rate for the 2SLS estimates is 0.08 for the specification with country–year fixed effects (a 10% increase in the PTR decreases employment probability by 0.8%). Both the OLS and 2SLS results are very robust to the inclusion of country–year fixed effects, which allow for country–specific changes in preferences for work as well as country–level changes in policies. Estimated coefficients of the control variables have the expected signs (see Appendix Table 2.7).

2.4.2 Heterogeneity in the responsiveness of labor supply

This section investigates the heterogeneity in responsiveness across groups of women by their education, family composition, and by groups of countries. I report how the 2SLS estimates of the PTR effect on female employment differ by woman's education in Panel A of Table 2.4. The effect of the PTR is largest for women with secondary education, for whom the 10 percentage point increase in the PTR decreases employment probability by 4.3 percentage points (the corresponding elasticity is 0.16). Both women with primary and tertiary education seem to be much less responsive to tax changes (an elasticity of 0.04 for both groups), but the effect is not significantly different from zero for women with primary education (possibly due to the small sample size). The non–linear relationship between the education and labor supply elasticity might be caused by a combination of two factors. On the one hand, less educated women have a lower employment rate (the means of the dependent variable can be found in the last column of Table 2.4), which usually implies larger employment elasticity. On the other hand, women with primary education tend to live in households with very low incomes, so that their participation on the labor market might be a necessity.

Panel B of Table 2.4 illustrates the differences in the responsiveness of female labor supply according to a woman's marital status and family composition. Married and cohabiting women are overall more responsive to tax and transfer changes than single women, which is a finding consistent with most previous studies (see, e.g. Bargain, Orsini, and Peichl 2014). The estimated PTR elasticity is 0.08 for married and cohabiting women and 0.06 for single women. However, when splitting the two groups by the presence of small children up to 5 years of age, a slightly different picture emerges. Single women with small children have by far the highest participation elasticity of 0.32, which is consistent with previous findings, where single mothers are the demographic group with the highest

Table 2.4: Estimates of the participation equation: heterogeneity of responses

	Dependent var.: employment dummy			
	2SLS			
	PTR (coeff.)	PTR (elast.)	N	dep. mean
Panel A. The PTR effect by woman's education				
Primary education	-0.093	-0.043	25,322	0.528
	(0.104)			
Secondary education	-0.429***	-0.161	250,280	0.796
	(0.041)			
Tertiary education	-0.125***	-0.043	158,005	0.919
	(0.033)			
Panel B. The PTR effect by family structure				
Married and cohabiting	-0.219***	-0.083	330,216	0.799
	(0.037)			
- with small children	-0.366***	-0.160	74,787	0.694
	(0.071)			
- without small children	-0.176***	-0.064	255,429	0.830
	(0.042)			
Single	-0.180***	-0.059	103,391	0.909
	(0.037)			
- with small children	-0.748***	-0.317	7,909	0.712
	(0.151)			
- without small children	-0.143***	-0.046	95,482	0.926
	(0.037)			
Panel C. The PTR effect by welfare regime				
Social–Democratic	-0.020	-0.007	38,011	0.964
	(0.052)			
Liberal	-0.361***	-0.144	21,744	0.794
	(0.061)			
Conservative–Corporatist	-0.121***	-0.052	97,883	0.854
	(0.046)			
Southern–European	-0.328***	-0.095	109,592	0.708
	(0.067)			
Post–Communist	-0.277***	-0.108	129,495	0.847
	(0.052)			
Former–USSR	-0.132	-0.038	36,882	0.898
	(0.261)			

Notes: The table reports the PTR coefficients from the 2SLS estimation applied separately to each group of women, the corresponding elasticity of employment to the PTR at the mean of independent variables, number of observations (N), and a mean of the dependent variable for each group. All regressions include a full set of country–year fixed effects and control variables. Standard errors in parentheses are clustered at country–year–group level (* $p<0.10$, ** $p<0.05$, *** $p<0.01$). The division of countries into welfare regimes is as following: DK, FI, SE belong to the social–democratic; IE, UK belong to the liberal; AT, BE, DE, FR, LU, NL belong to the conservative–corporatist; CY, GR, IT, PT, ES belong to the Southern–European; BU, CZ, HU, PL, RO, SI, SK belong to the post–communist; and EE, LT, LV belong to the former–USSR welfare regime. Source: EUROMOD model and EU-SILC data (2005–2010), own calculations.

participation elasticity (Meghir and Phillips 2008). Married and cohabiting women with small children also have a very large participation elasticity of 0.16, while both single and coupled women without small children have a substantially lower responsiveness towards tax and transfer changes (corresponding elasticities of 0.06 and 0.05 for coupled and single women, respectively).

Finally, I investigate the heterogeneity of the effect of the PTR on employment by groups of countries. For this purpose, I use a well–known welfare regime typology by Esping-Andersen (1990) that creates groups of countries based on social policies and organization of work. Esping-Andersen (1990) differentiated between three models of the welfare state: the social–democratic, the liberal, and the conservative–corporatist welfare state. This typology was later extended with the Southern-European welfare regime (Ferrera 1996), and the European post–communist and former–USSR categories (Fenger 2007). I use this extended welfare regime categorization of Fenger (2007), which allows the categorizing of all countries in the sample.[23]

Panel C of Table 2.4 reports the PTR coefficients and elasticities by welfare regime. The participation elasticities are highest in the liberal (0.14), the post–communist (0.11), and the Southern–European (0.10) welfare regimes. These are also the three groups of countries with the lowest female employment. Therefore, the results are consistent with previous findings that elasticities are larger in countries where female employment is lower (Blau and Kahn 2007, Bargain, Orsini, and Peichl 2014). The responsiveness is substantially lower in the conservative–corporatist welfare regime, and not significantly different from zero in the social–democratic and the former–USSR welfare regimes. This is not very surprising given that the female employment rates in these two groups of countries reach 96.4% and 89.8%, respectively.

Results presented in this section are also in line with Bargain, Orsini, and Peichl (2014), which is the closest paper to this study. Due to methodological differences and the fact that I estimate elasticities with respect to participation tax rates, while Bargain, Orsini, and Peichl (2014) estimate responsiveness to net wages and non–labor income, a direct comparison of my results with Bargain, Orsini, and Peichl (2014) is rather difficult. Nevertheless, both studies have found substantially smaller female labor supply elasticities than was found in most of the previous literature (Arellano and Meghir 1992, Laroque and Salanié 2002, van Soest, Das, and Gong 2002, Callan, van Soest, and Walsh 2009). Both studies also found a substantial heterogeneity of labor supply elasticities across coun-

[23] See the note below Table 2.4 for the division of countries into welfare regimes.

tries (with the UK and Southern–European countries being among the largest–elasticity countries), and across groups of women (with the largest participation elasticity among single mothers).

2.4.3 Examining the quality of the PTR imputation

In this section, the quality of the PTR imputation to the EU–SILC data is examined (for details on the imputation, see Section 2.3.1). To illustrate the impact of the PTR imputation on the estimated coefficients, I compare the results of the participation equation based on the underlying EUROMOD data and on the EU-SILC data (with the imputed PTR) for the same set of country–year cells. The main difference between the two datasets is that the underlying EUROMOD data include the calculated PTR, while the EU–SILC data only have the imputed PTR. I restrict the sample to all countries for which the EUROMOD data are available for 2005–2007.

The comparison of results is provided in Appendix Table 2.6. The PTR is highly significant in the 2SLS specification and the magnitude of coefficients is quite similar using both the EUROMOD data and the EU–SILC data with the imputed PTR—the coefficients are -0.52 based on the EUROMOD data and -0.44 based on the EU–SILC in the specification with country–year fixed effects. The elasticities are also quite similar, although again somewhat larger in the estimation using EUROMOD data, with elasticity of 0.18 as compared to 0.15 based on the EU–SILC.[24]

The estimated coefficients in the OLS specification are quite different using the EUROMOD data and EU–SILC data. In fact, the results based on the EUROMOD data suggest a positive effect of the PTR on employment probability. Therefore, the magnitude of bias in the OLS estimation seems to be much higher than is suggested by the EU–SILC estimation results (see Appendix Table 2.6). Overall, the estimation results based on the EU–SILC data with the imputed PTR thus seem to be somewhat smaller in magnitude than those based on the EUROMOD data, suggesting the presence of a downward bias caused by the imputation procedure.[25] Nevertheless, the main findings

[24]The estimated coefficients and elasticities presented here are much larger in magnitude than the estimates presented in Section 2.4.1. This is because they are based on a selected group of counties for which the EUROMOD data are available for years 2005–2007. This sample selection was mainly driven by the necessity to have a sufficient number of tax and transfer changes in the sample to be able to identify the effect of the PTR on employment decisions in the EUROMOD data.

[25]As the imputation procedure increases measurement error in the PTR, the downward bias was to be expected.

remain unchanged whether we use the EUROMOD data or the data with the imputed PTR.

2.5 Conclusion

This paper investigates the impact of tax and transfer policies in the countries of the EU on the extensive margin of female labor supply. Unlike previous studies, I utilize an indicator of extensive margin work incentives—the participation tax rate—as the main explanatory variable. This allows me to capture of the effect of both tax and benefit systems on the work incentives of women and to deal with possible endogeneity of the participation tax rate by using a simulated instrumental variable. The instrumental variable allows me to exploit only the variation in the participation tax rate due to changes in policies, setting aside the variation due to measurement error and endogenous behavioral responses. Further, the rich structure of the data, which cover 26 EU countries from 2005–2010, allows me to control for time–invariant and time–varying country–level unobserved factors (such as economic shocks, tastes for work, or family policies), and use all policy changes that took place in these countries between 2005 and 2010 for identification.

The results suggest that the labor participation decisions of women in the EU are indeed negatively affected by the level of effective taxation they face. The comparison of estimates based on the OLS and the IV approaches confirms the presence of attenuation bias caused by the measurement error and income endogeneity. The instrumental variable estimation implies that a 10 percentage point increase in the participation tax rate decreases the female employment probability by 2 percentage points (the corresponding participation elasticity is 0.08). The effect is higher for secondary educated women—results suggest that they respond to the 10 percentage point increase in the PTR by decreasing their employment probability by 4.3 percentage points (elasticity of 0.16). Women with primary and tertiary education are substantially less responsive to tax and transfer incentives (elasticity of 0.04 for both groups).

I also investigate the heterogeneity of responses towards tax and transfer systems across groups of women by their marital status and family composition. The results are in line with previous findings (Keane 2011, Bargain, Orsini, and Peichl 2014). The responsiveness is higher for married women than for single women, for women with small children, and the highest elasticities are found for single mothers (participation elasticity of 0.32).

Further, I use the typology of welfare regimes originally proposed by Esping-Andersen (1990) to uncover the heterogeneity of responses across groups of countries. The results indicate that the effect of the PTR on employment probability is the highest for the liberal welfare regime (to which Ireland and the UK belong), the post–communist and the Southern–European welfare regimes. These are also the three groups of countries with the lowest female employment rate. Therefore, the findings are consistent with previous findings that suggest that the higher the female labor force participation, the lower the female labor supply elasticity (Blau and Kahn 2007).

Bargain, Orsini, and Peichl (2014) is, to my knowledge, the only study that uses a multinational tax–benefit microsimulation model in the analysis of female labor supply behavior, and is thus closest to the present study. Similarly, I find much smaller female participation elasticities than some previous studies, and a substantial heterogeneity of elasticities across groups of women and across countries. In particular, single mothers and women living in Southern Europe and the UK have the largest participation elasticities, according to both the present study and Bargain, Orsini, and Peichl (2014). However, the present study takes a different estimation approach and uses a much longer time span than Bargain, Orsini, and Peichl (2014). My sample also covers 7 post–communist and 3 former–USSR countries, which have been only rarely studied in the female labor supply literature, while Bargain, Orsini, and Peichl (2014) cover only Estonia, Hungary and Poland from this group. My results suggest that the post–communist countries have among the highest female participation elasticities, while the former–USSR countries belong to the group of countries with the highest female employment and lowest participation elasticities.

Multinational microsimulation models offer a very useful tool for the study of tax and transfer impact on the labor supply, as they allow for large scale international comparison as well as a comprehensive analysis of heterogeneity in responsiveness across groups of individuals. However, they have been rarely used in the labor supply elasticity literature. The main shortcoming of these models is the relatively short time span that they cover to date. Therefore, future research should take advantage of a much richer time variation in policies that will be available in the multinational microsimulation models and use it to further reconcile the persistent controversy over the responsiveness of labor supply to tax and transfer changes.

2.A Appendix 2

Table 2.5: Overview of country–year cells used in the analysis

	Tax–benefit rules						Input data		
	2005	2006	2007	2008	2009	2010	2005	2006	2007
AT			x	x	x	x			x
BE	x	x	x	x	x	x	x	x	x
BG			x	x	x	x			x
CY		x	x	x				x	x
CZ	x	x	x	x	x	x	x	x	x
DE			x	x	x	x			x
DK			x	x	x	x			x
EE	x	x	x	x	x	x	x	x	x
FI			x	x	x	x			x
FR		x	x	x	x	x		x	
GR	x	x	x	x	x	x	x	x	x
ES	x	x	x	x	x	x	x	x	x
HU		x	x	x	x	x		x	x
IE		x	x	x				x	x
IT	x	x	x	x	x	x	x	x	x
LT	x	x	x	x	x	x	x		x
LU			x	x	x	x			x
LV		x	x	x	x	x		x	x
NL		x	x	x	x	x		x	x
PL		x	x	x	x	x		x	x
PT		x	x	x	x	x		x	x
RO			x	x	x	x			x
SE		x	x	x	x	x		x	x
SI		x	x	x	x	x		x	x
SK		x	x	x	x	x		x	x
UK		x	x	x	x	x			x

Notes: Table illustrates for which country–year cells the EUROMOD tax–benefit computations are available and for which cells the EUROMOD input data are available. The main results are based on all country–year cells, for which the tax–benefit rules are available.

Table 2.6: Comparison of results based on EUROMOD data and EU-SILC data with the imputed PTR

	(1)	(2)	(3)	(4)
	Dependent var.: employment dummy			
	OLS		2SLS	
Panel A: Results based on EUROMOD data (2005-2007)				
PTR (coefficient)	0.330***	0.329***	-0.434*	-0.519**
	(0.090)	(0.090)	(0.250)	(0.242)
PTR (elasticity)	0.115	0.115	-0.151	-0.181
country–year fixed effects	no	yes	no	yes
R^2	0.201	0.201	0.179	0.174
Observations	75,928	75,928	75,928	75,928
Panel B: Results based on EU-SILC data (2005-2007)				
PTR (coefficient)	-0.048*	-0.056**	-0.361***	-0.438***
	(0.027)	(0.025)	(0.137)	(0.081)
PTR (elasticity)	-0.016	-0.019	-0.121	-0.147
country–year fixed effects	no	yes	no	yes
R^2	0.181	0.182	0.177	0.175
Observations	75,066	75,066	75,066	75,066

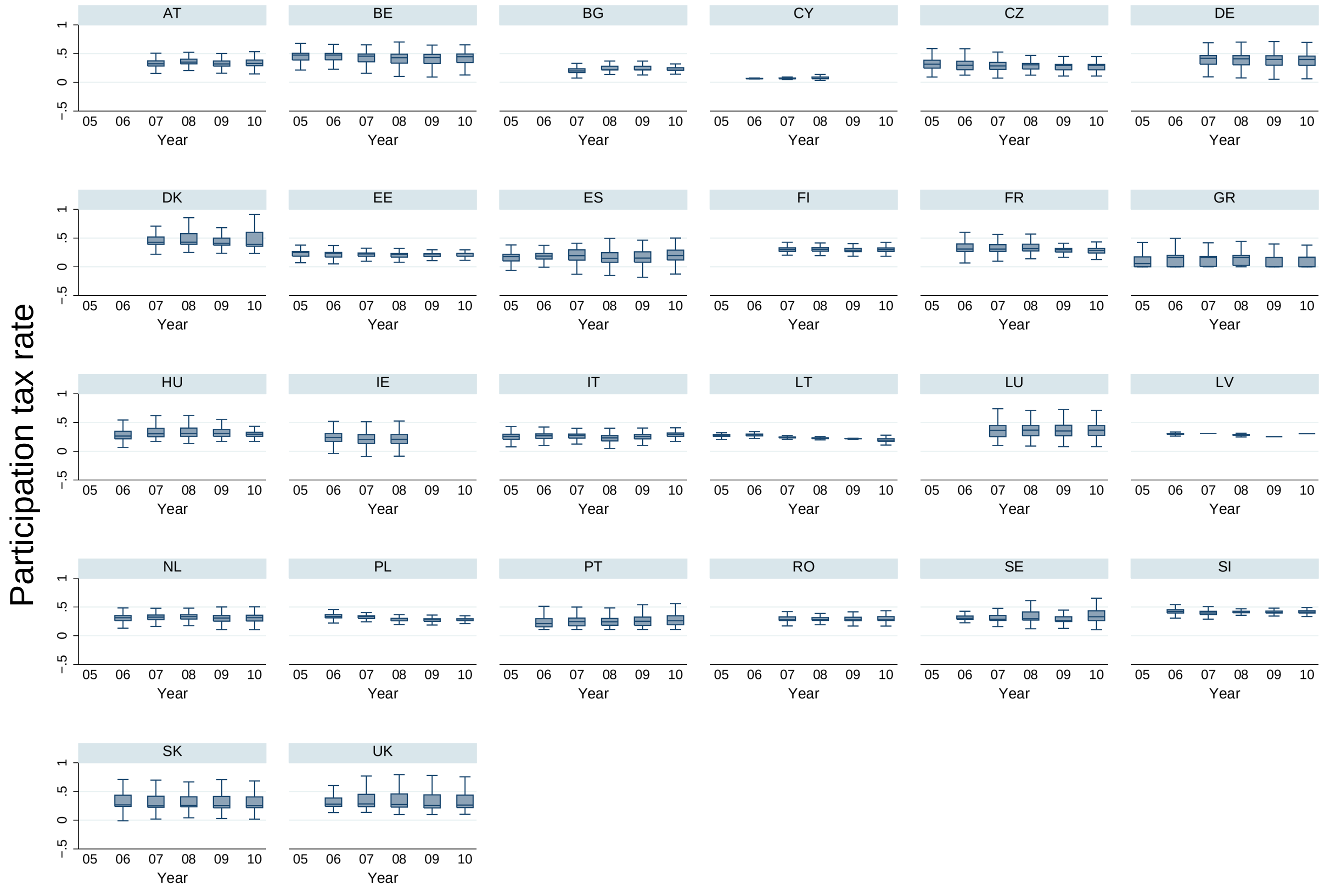
Notes: All regressions include a full set of country dummies, year dummies, and control variables. The sample is restricted to the same set of country–year cells for both samples—to all countries, which have years 2005–2007 covered in the EUROMOD input data. Standard errors in parentheses are clustered at country–year–group level, where groups are defined by education, presence of children of various ages, and marital status (* p<0.10, ** p<0.05, *** p<0.01).
Source: EUROMOD and EU–SILC data (2005–2007), own calculations.

Table 2.7: Estimates of the participation equation, full specification

	(1)	(2)	(3)	(4)
	Dependent var.: employment dummy			
	OLS		2SLS	
PTR	-0.060***	-0.061***	-0.188***	-0.208***
	(0.009)	(0.008)	(0.039)	(0.030)
secondary education	0.201***	0.200***	0.202***	0.201***
	(0.009)	(0.009)	(0.009)	(0.009)
tertiary education	0.302***	0.302***	0.306***	0.306***
	(0.008)	(0.008)	(0.008)	(0.008)
age	0.030***	0.030***	0.030***	0.030***
	(0.001)	(0.001)	(0.001)	(0.001)
age squared	-0.000***	-0.000***	-0.000***	-0.000***
	(0.000)	(0.000)	(0.000)	(0.000)
child aged 1	-0.090***	-0.090***	-0.090***	-0.090***
	(0.006)	(0.006)	(0.006)	(0.006)
child aged 2	-0.222***	-0.222***	-0.222***	-0.222***
	(0.013)	(0.013)	(0.013)	(0.013)
child aged 3	-0.171***	-0.171***	-0.170***	-0.170***
	(0.011)	(0.011)	(0.011)	(0.011)
child aged 4-5	-0.074***	-0.074***	-0.074***	-0.074***
	(0.005)	(0.005)	(0.005)	(0.005)
child aged 6-9	-0.056***	-0.056***	-0.056***	-0.055***
	(0.004)	(0.004)	(0.004)	(0.004)
child aged 10-15	-0.036***	-0.037***	-0.036***	-0.036***
	(0.004)	(0.004)	(0.004)	(0.004)
married	-0.101***	-0.101***	-0.102***	-0.101***
	(0.006)	(0.006)	(0.006)	(0.006)
cohabiting	-0.055***	-0.055***	-0.056***	-0.056***
	(0.005)	(0.005)	(0.005)	(0.005)
number of HH members	-0.021***	-0.021***	-0.022***	-0.022***
	(0.001)	(0.001)	(0.001)	(0.001)
presence of elderly in the HH	-0.003	-0.003	-0.003	-0.003
	(0.003)	(0.003)	(0.003)	(0.003)
inactive partner	-0.042***	-0.042***	-0.042***	-0.042***
	(0.003)	(0.003)	(0.003)	(0.003)
secondary–educated partner	0.042***	0.042***	0.042***	0.042***
	(0.004)	(0.004)	(0.004)	(0.004)
tertiary–educated partner	0.058***	0.058***	0.059***	0.058***
	(0.004)	(0.004)	(0.004)	(0.004)
country–year fixed effects	no	yes	no	yes
R^2	0.174	0.175	0.173	0.173
Observations	433,607	433,607	433,607	433,607

Notes: All regressions include a full set of country and year dummies. Standard errors in parentheses are clustered at country–year–group level, where groups are defined by education, presence of children of various ages, and marital status (* p<0.10, ** p<0.05, *** p<0.01). Source: EUROMOD model and EU-SILC data (2005–2010), own calculations.

Figure 2.1: Evolution of the PTR for women in the EU countries between 2005 and 2010



Notes: Each box plot indicates the 25th and 75th quartile (the bottom and top of the box), the median (the band inside the box), and the lowest (highest) datum still within 1.5 interquartile range of the lower (upper) quartile (the ends of the whiskers).
Source: EUROMOD model and EU–SILC data (2005–2010), own calculations.

Chapter 3

Career Breaks after Childbirth and the Role of Parental Leave Policies

Co-authored by Alena Bičáková

3.1 Introduction

Career development and remuneration for work grow with work experience and job tenure. Despite increasing participation of fathers in caring for their children, career breaks around childbirth are still one of the major sources of gender differences in labor market outcomes.[1]

The total length of the career breaks around childbirth is affected by preferences, family leave policies, and childcare availability, as well as labor market conditions. Family leave of several months allows mothers to bond with their child, while keeping their jobs, and thus increases female labor force participation, employment and future earnings (Hashimoto et al. 2004, Baker and Milligan 2008, Han, Ruhm, and Waldfogel 2009). Family leave of several years per child, on the other hand, represents a major break in one's career that is likely to result in human capital deterioration and a decrease in productivity, thus having an opposite effect on female labor market outcomes (Schone

[1] Negative effects of career breaks on women's wages are documented e.g. by Anderson, Binder, and Krause (2002), Spivey (2005), Miller (2011), Ejrnaes and Kunze (2013). Impact on other labor market outcomes, such as loss of human capital or occupational choice is studied e.g. in Francesconi (2002), Adda, Dustmann, and Stevens (2011). Previous studies have documented a strong positive correlation between duration of paid parental leave and gender wage gaps (OECD 2012), as well as between paid parental leave duration and gender unemployment gaps (Bičáková 2012).

2004, Lalive and Zweimüller 2009, Schönberg and Ludsteck 2014). Therefore, the impact of changes in the statutory family leave on mothers' post–birth employment depends on its duration, the size of the allowance, its take-up rate, job protection provisions (Schönberg and Ludsteck 2014), and childcare availability, as well as on social norms related to family gender roles and on the overall attachment of women to the labor market (Bergemann and Riphahn 2015).

This paper considers the impact of the duration of paid parental leave on the overall post–birth career interruptions of mothers in the Czech Republic, using two reforms of parental leave allowance, in 1995 and 2008, respectively. The 1995 reform increased duration of paid parental leave from three to four years. The 2008 reform reduced it for some women to three or two years and made the duration partly choice–based. Although the two reforms altered the duration of receipt of the allowance, the job protection regulations that entitle women to return to their pre–childbirth jobs remained unchanged, at three years post birth. We therefore focus on the impact of the changes in the paid parental leave solely through its monetary incentives, unaccompanied by changes in job security.[2]

We argue that the Czech Republic, and the two reforms, represent a unique economic context that helps us bring new evidence about the impact of paid parental leave on total career interruptions after childbirth: Czech women have very high female labor force participation (exceeding 80%), a heritage of the Communist regime (see Fodor 2005). This strong overall attachment of Czech women to the labor market contrasts sharply with the absence of mothers of children younger than 3 in the labor market: The Czech Republic has one of the longest paid parental leaves with one of the highest take up rates in the EU (OECD 2010). Given the lack of almost any institutional childcare for children younger than three,[3] the full duration of the parental leave is used by the majority of mothers.[4] Such a constellation places the Czech Republic among the three EU countries

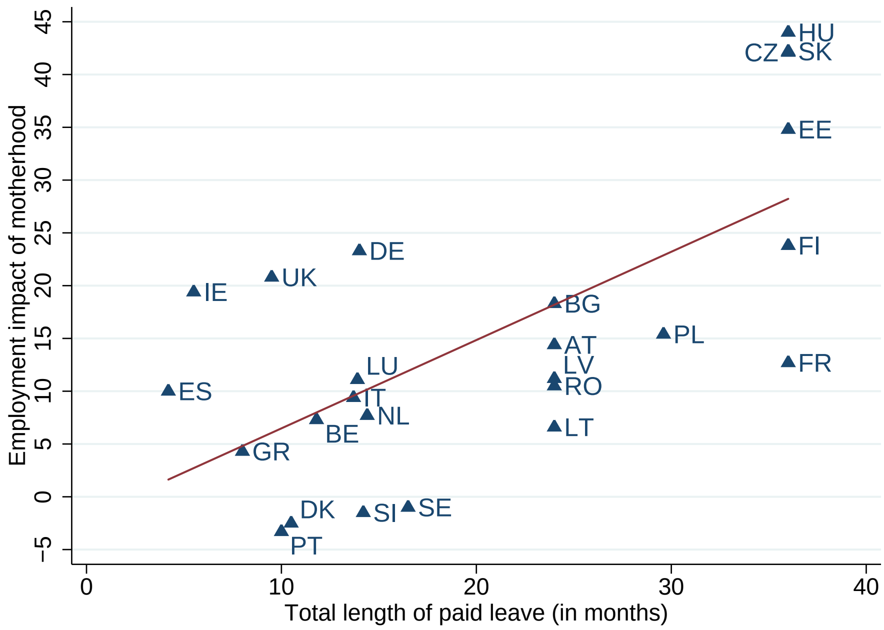
[2]The majority of previous studies regarding the impact of paid parental leave cannot disentangle the two effects, as the duration of the benefit and that of job protection are identical in most countries, and reforms alter both in the same way. The only exceptions were the 1992, 1993, and 2007 German reforms studied in Schönberg and Ludsteck (2014) and Bergemann and Riphahn (2015). The Czech reforms differ from these by a longer parental leave duration and much higher take up rate (80% in the Czech Republic versus 65% in Germany, see OECD 2010).

[3]Only 0.4% of children younger than 3 were covered by institutional childcare in 2012 (source: Institute for Health Information and Statistics of the Czech Republic).

[4]When we exclude mothers with children between 0 and 3, the labor force participation of the prime–age Czech women is as much as 93%. The situation is similar in other Central and Eastern European countries, but the employment rate of women without small children is not as high as in the Czech Republic (Fodor 2005).

(together with Slovakia and Hungary) with the the most sizable consequences of childbirth on mothers' employment. Figure 3.1 highlights the strong positive correlation between the duration of paid family leave and the impact of employment on motherhood, defined as a percentage point difference between the employment rate of childless women and women with at least one child below 6 years of age.

Figure 3.1: Family leave and employment impact of motherhood in the EU countries.



Note: The total length of paid leave is the statutory maximum length of postnatal paid leave (in months) in 2012. The employment impact of motherhood is the difference between the employment rate of childless women and women with at least one child below 6 years of age (in percentage points) in 2013. Source: Annual reviews of the International Network on Leave Policies and Research and EUROMOD country reports for the length of leave. Eurostat: Employment rate of adults by sex, age groups, highest level of education attained, number of children and age of youngest child for the employment impact of motherhood.

Our data reveal that the initial period of a post–birth family leave (phase of inactivity) is often followed by a spell of unemployment.[5] Some mothers may not be entitled to the three–year job protection and some may extend their family leave beyond the protected period. There is also informal evidence questioning the strength and enforceability of the job protection especially in the transition economies and in times of economic downturn (Kantorova 2004, Fodor 2005). The data suggest that over 80% of Czech mothers who are unemployed before their child turns 5 enter unemployment immediately after the end

[5] 10% of Czech mothers with a youngest child aged 3 were unemployed in 2012 and the share exceeded 13% for those with children aged 4 and 5 (source: Czech LFS data, own calculations).

of the family leave. Even when women keep their pre–childbirth jobs, the unemployment spell may follow soon after they return from the leave.[6]

Despite such evidence, previous research on the impact of family leave policies on labor market status of women does not distinguish between family leave and subsequent unemployment. All studies we are aware of focus solely on mothers' post–childbirth employment, i.e. the overall career break. The only paper that explores the impact of family leave policies on unemployment is Das and Polachek (2014). However, they consider the unintended effects of changes in family leave regulations on unemployment of all young women, and not specifically that of mothers after childbirth. This study aims to fill the gap. After estimating the impact of the two reforms of parental allowance on non–employment, i.e. the total career break after childbirth, we focus on post–birth inactivity and unemployment separately, using the International Labor Organization definition of unemployment, based on the self–reported information provided in the data.[7]

While there is a strand of literature that questions the difference between the two labor market states (Flinn and Heckman 1983, Gönül 1992, Benati 2001) and the usefulness of the ILO definition Jones and Riddell 1999, Brandolini, Cipollone, and Viviano 2006), our study suggests that career breaks after childbirth represent a situation where the distinction between the two states is both meaningful and important. Moreover, a spell of unemployment after family leave is likely to have a negative impact both on women's contemporaneous utility as well as on the quality of post–childbirth jobs (Francesconi 2002, Adda, Dustmann, and Stevens 2011, or result in permanent withdrawal from the labor force (the so–called 'discouraged worker effect').

In sum, we extend the previous literature along several dimensions: we estimate the impact of changes in paid parental leave longer than three years, we explore the impact of two reforms that altered the duration of parental allowance in the opposite direction, while keeping job protection the same, and we focus on a country with a very high take–up rate of the parental leave, but also strong overall attachment of women to the labor market. Most importantly, having estimated the impact of the reforms on total career interruptions due to childbirth, we then disentangle the post–childbirth non–employment of mothers into inactivity and unemployment, thus reflecting the distinction between the actual duration of family leave and the period of job search that often follows immediately

[6]Schönberg and Ludsteck (2014) report for Germany that 7% of women who return to work at the end of the statutory leave become unemployed 1 or 2 months afterwards.

[7]We classify an individual as unemployed if he/she does not have a job, is actively seeking a job, and is ready to start working within two weeks.

or soon after the end of the leave.

We estimate the impact of the reforms on labor market status of women separately by the age of the youngest child, using a standard differences-in-differences design. While data limitations and the nature of the reforms do not allow us to estimate the effect of changes in duration of paid parental leave on career breaks with a regression discontinuity design around the birth of the child, we follow the standard approach in the literature and use women with older children to control for the aggregate trends and business cycle developments in the labor market.[8]

A recent study by Müllerova (2014) does apply a regression discontinuity approach to estimate the impact of the 1995 reform on mothers' employment in the Czech Republic using the same data, but it does not focus on the 2008 reform and—in particular—does not differentiate between inactivity and unemployment. The small samples of mothers with a child born around the date of the reforms, as used in the regression discontinuity approach, would not allow us to estimate the effect by the age of the youngest child on inactivity and unemployment separately.[9] The relevant subset of our results is similar to Mullerova's findings, suggesting that our identification strategy should be sound.

We find that the 1995 reform prolonged the parental leave of at least one third of mothers and shifted the post–leave unemployment spell to the time when a child turns 4. The 2008 reform, on the other hand, shortened the parental leave of at least one fifth of mothers and shifted the post–leave job search to the time when a child turns 2 or 3. The two reforms have not affected mothers of children older than 7 or 6 respectively, suggesting that the consequences of the changes in the paid parental leave duration on the labor market status of women after childbirth are only medium term. Future research is needed to explore the potential long–term consequences of the overall duration of family career breaks on income and quality of the jobs that women hold after childbirth.

The paper is organized as follows: The following section is devoted to the institutional and demographic background of the two reforms. We then present our estimation strategy and descriptive evidence. The results section is followed by a robustness analysis and the conclusion.

[8] A similar approach was used, for example, in Naz (2004), Schone (2004), Sánchez-Mangas and Sánchez-Marcos (2008), Geyer, Haan, and Wrohlich (2014), Bergemann and Riphahn (2015).

[9] This and the fact that the information about a child's age is imprecise in the data were the main reasons we decided to use the stronger identification assumptions required for the difference–in-differences approach, but richer data that allowed us to address a wider range of questions.

3.2 Institutional background

3.2.1 Family leave policies

Family leave policies in the Czech Republic consist of three components: job protection, maternity benefit, and parental allowance. Czech parents are eligible for *job–protected leave* until the child's third birthday.[10] The three–year long job protection period was introduced in 1990 and is still in force nowadays. However, the strength of job protection in transition economies has been often questioned in the literature due to an unstable business environment, in which companies disappear quickly, and often rationalize or cancel branches and work positions (Kantorova 2004, Fodor 2005). Moreover, there is some evidence that Czech employers avoid job protection and show unwillingness to reemploy women on parental leave even in times of economic prosperity (Kucharova 2006).

Czech women who were employed for at least 270 days in two years prior to a child's birth are entitled to receive *maternity benefit* for 28 weeks (starting 6 to 8 weeks prior to birth). From 1990 to 2007, the maternity benefit paid 69% of a woman's salary in the last 12 months prior to the commencement of maternity leave. In 2008, it was increased by one percentage point to 70% of a woman's salary. There have been no other substantial changes to maternity benefit.[11]

A parent taking care of a child is also eligible for *parental allowance*, a non–means–tested flat rate benefit. Receipt of parental allowance starts either immediately after the end of maternity benefit or right after childbirth if the mother is not eligible for maternity benefit. The eligibility criteria for parental allowance required the parent not to earn above a certain threshold (in force until 2004) and the child not to attend a childcare facility for more than a certain amount of time per month (in force until nowadays).[12]

[10] The eligibility requirements are relatively modest. All employees employed on a permanent contract are eligible for job protection. An employee with a fixed–term contract is also eligible for job protection, but only up to the date of contract expiration.

[11] The amount of maternity benefit is reduced for higher income levels using reduction bands and has a maximum threshold. The reduction bands have increased gradually over time since the early 1990s. There was no substantial change in the reduction bands with the exception of 2009, when the maximum daily amount of maternity leave increased from CZK 479 to 963 per day. However, this change only affected benefits for women with very high wages (the maximum amount applied to women with a gross monthly income of CZK 71,000, which was more than three times the average female wage at that time).

[12] The maximum threshold for earnings increased over time throughout the 1990s and was completely abolished in 2004. However, the threshold on hours per day and days per week that a child can attend a childcare facility is still in place. Therefore, if a parent wanted to collect the allowance and work at the same time, the grandparents would need to take care of a child or a baby–sitter would have to be hired.

The 1995 reform

Until 1995, the maximum length of receipt of parental allowance coincided with the length of job–protected leave, i.e. a parent could collect the allowance until the child's third birthday when the job protection ended. In October 1995, a reform was introduced that prolonged the receipt of parental allowance until the child's fourth birthday, while the length of job–protected leave remained unchanged. The benefit duration, therefore, suddenly exceeded the period of job protection by 12 months. All parents with children under 4 years of age as of October 1, 1995 were eligible for the prolonged parental allowance. The reform thus affected not only the parents of newborn children, but also those with older children who met the eligibility criteria.

The monthly allowance payment was kept almost unchanged in 1995—it increased only slightly from CZK 1,740 to CZK 1,848 (the allowance corresponded to about one fourth of an average wage in the economy). This minor change was part of the gradual increase in the monthly allowance over the 1990s and early 2000s due to the rise in the minimum subsistence income it was derived from (Figure 3.2). When evaluating the impact of the 1995 reform, we focus on 1993–1999, during which gradual increases in the monthly benefit amount experienced only negligible changes relative to the extension of the duration of the allowance by one year, and the corresponding rise in the total amount of benefit received.

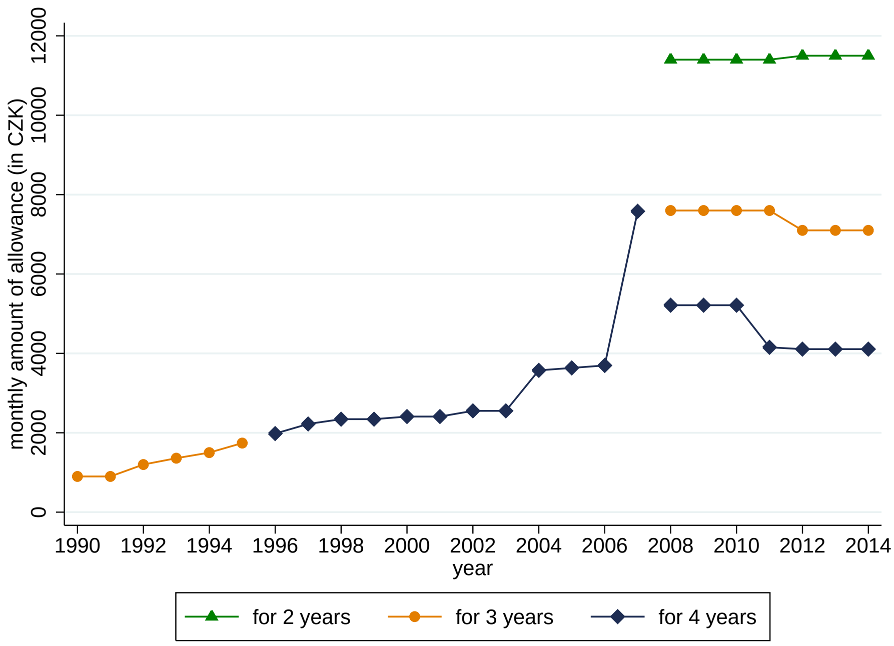
A substantial increase in the amount of the parental allowance took place only in 2007, when the monthly amount doubled (from CZK 3,696 to CZK 7,580). However, this increase was immediately followed by an important reform of the parental allowance schedule in 2008.

The 2008 reform

Since January 1, 2008, parents have been allowed to choose the length and the corresponding level of monthly parental allowance. The shortest track paid CZK 11,400 per month until the child's second birthday, the standard track paid CZK 7,600 until the child's third birthday, and the longest track paid CZK 7,600 until the child was 21 months old and then CZK 3,800 until the child's fourth birthday (see Figure 3.2). All parents were entitled to the four–year track. Entitlement to the three–year track was conditional on

Given the very low share of part–time jobs in the Czech Republic, it was very difficult for most parents to combine receipt of parental allowance and a formal job in the 1990s, and this is still true today.

Figure 3.2: Monthly parental allowance in the Czech Republic, 1990-2014.



Note: The figure depicts the average monthly amount of parental allowance for each maximum length of receipt of the allowance.

one of the parents working for at least 270 days in the two years prior to birth. If, in addition, one parent earned on average at least CZK 16,500 per month in the 12 months prior to the birth,[13] they were also eligible for the two–year track.

The new system of parental allowance covered not only parents of children born after January 2008, but also to some extent those born before. In particular, all parents of children younger than 22 weeks as of January 2008 could choose the two–year, three–year or four–year track, and parents with children younger than 21 months as of January 2008 were eligible for the three– or four–year track, conditional on fulfilling the other eligibility criteria. Therefore, women who took advantage of the two–year track of parental allowance might have returned to the labor market after two years of leave no sooner than in August 2009, and women who chose the three–year track might have returned to the labor market after three years of leave no sooner than in May 2009.

There were some minor changes to the parental allowance scheme after 2008. The monthly amounts of the allowance decreased for the four–year track in 2011 (it only paid CZK 7,600 until the child was 9 months old and then CZK 3,800 until the child's fourth birthday) and for the three–year track in 2012 (from CZK 7,600 to CZK 7,100 per month).

[13]In 2008, an average male wage was CZK 29,429 and an average female wage CZK 21,789.

The purpose of these minor changes was to unify the total amount of allowance per child for all tracks. Since 2012, the maximum total amount of allowance per child has been CZK 220,000, regardless of the length of receipt of parental allowance.

When we evaluate the impact of the 2008 reform, we focus on 2004–2012. We abstract from the other changes, as they are minor relative to the introduction of the flexible parental allowance system and they do not alter the duration of paid parental leave.[14]

3.2.2 Unemployment benefits

As we focus on the labor market status of mothers with young children, and strive to disentangle the post–birth family leave (inactivity) from the potential post–leave unemployment spell, we also have to consider the monetary incentives provided by the unemployment benefit scheme. In general, unemployed individuals in the Czech Republic are eligible for unemployment benefits during the first six months of unemployment if they worked for at least 12 months in the prior three years. During the 1990s, the benefit paid 60% of previous monthly earnings for the first three months and 50% of monthly earnings for the next three months of unemployment.[15]

An individual is also eligible for unemployment benefit if s/he took care of a child below 3 years of age for at least 12 months in the last three years. However, if the eligibility for unemployment benefit is based on the time spent taking care of a child instead of work, the amount of unemployment benefits is much lower, corresponding to about 55% of the monthly parental allowance after 1995 and about 34% of the allowance for the three–year track after 2008.[16] Therefore, most Czech women who become unemployed immediately after parental leave are eligible only for this substantially lower unemployment benefit level. Post–leave unemployment therefore represents a substantial drop in monthly funds available to mothers of young children, which is likely to decrease their reservation wage when searching for a job.

[14]The 2007 increase in the monthly parental allowance was quite substantial, but it was in place for one year only.

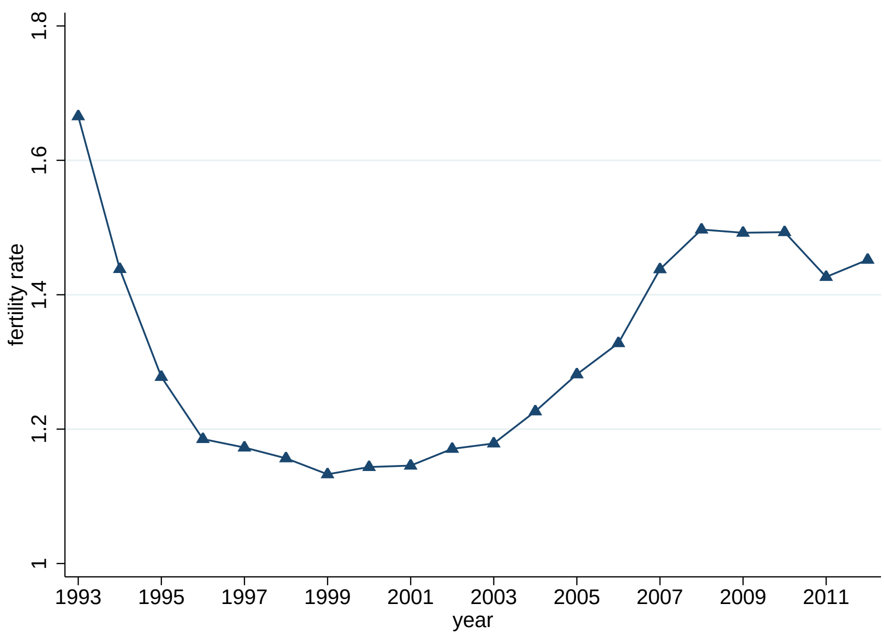
[15]In 1998, the amount went down slightly to 50 and 40% of previous earnings, and in 2005, it was again increased slightly to 50 and 45% of previous earnings, respectively. In 2009, the length of benefit receipt was shortened to five months, paying 65, 50, and 45% of previous earnings in the first 2 months, next 2 months, and the last month, respectively.

[16]The unemployment benefit is in this case calculated not from the previous wage, but from the minimum living standard, which corresponded to around 10–20% of the average wage in the economy throughout the 1990s and 2000s.

3.2.3 Fertility and childcare availability

The period of the 1990s was marked by a steep decline in fertility rates in most transition economies (Sobotka 2003), the Czech Republic being no exception (see Figure 3.3). The transformation to democracy and a market economy was accompanied by a decline in real wages, changes in the life styles of young couples and other factors that negatively affected fertility. The fertility rate of Czech women dropped from 1.67 in 1993 to 1.13 in 1999. Sobotka (2003) argues that in central Europe, this decline was mainly caused by postponement of parenthood (the average age of Czech women at first birth increased from 22.6 in 1993 to 24.6 years in 1999). The fertility rate then started to increase slowly in the early 2000s and reached its peak of 1.5 in 2008 (Figure 3.3). This increase was mainly a consequence of a generation of baby boomers from the 1970s entering childbearing age.

Figure 3.3: Fertility rates, 1993–2012

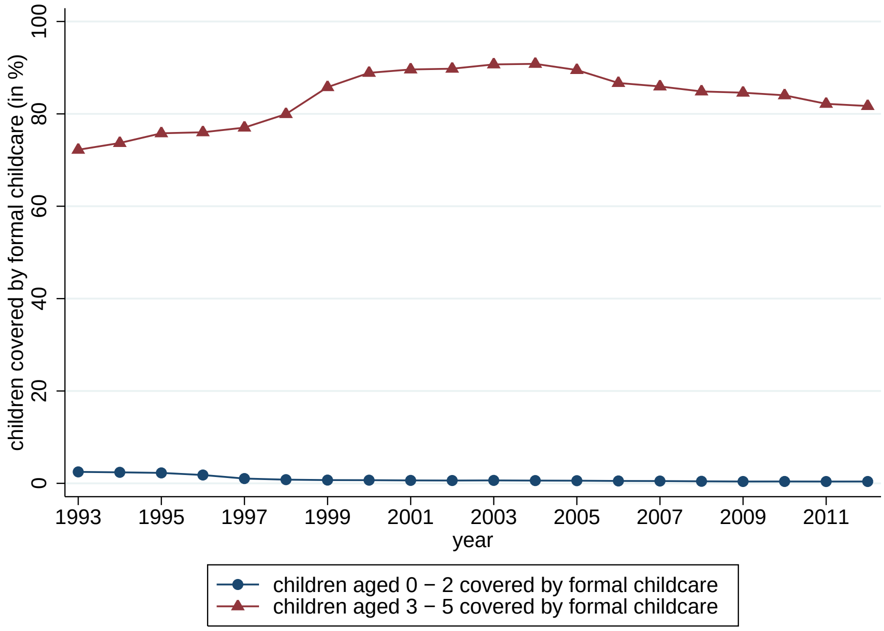


Note: The figure depicts fertility rates in the Czech Republic in 1993–2012. Source: Czech Statistical Office.

As for the availability of public childcare, pre–school education in the Czech Republic consists of nurseries (facilities for children below 3 years of age) and kindergartens (for children aged 3 and above). Nurseries are fairly scarce in the Czech Republic. Only around 0.4% of children below 3 attend public nurseries (see Figure 3.4) and private nurseries are also very scarce.[17] This is to a large extent a result of massive closures

[17]Official statistics for the coverage by private private nurseries are not available, but Eurostat esti-

Figure 3.4: Coverage of pre–school children by formal childcare



Note: The figure depicts formal childcare coverage for children aged 0–2 and 3–5 in the Czech Republic in 1993–2012. The coverage is calculated as a share of children of a given age attending public childcare to all children from this age group. Source: Institute for Health Information and Statistics of the Czech Republic and Czech Statistical Office.

of public nurseries in the early 1990s driven by a substantial drop in fertility, but also by a broader change in family policy, which largely promoted conservative policies that encouraged women to leave the labor market to raise their children (Saxonberg and Sirovátka 2006).[18]

The situation is markedly different for children between 3 and 5 (i.e., children below the compulsory school age of 6), as about 80% of them are enrolled in formal public childcare. The supply of kindergartens was also slightly reduced in the 1990s, but the decline was not substantial, so that the overall coverage of children between 3 and 5 years increased slightly in the 1990s as a consequence of a large fertility drop. The coverage started to diminish only in 2005 when the baby boom commenced (see Figure 3.4).

During the two periods we study, there were therefore no substantial changes in childcare availability. While fertility evolved rather dynamically over the same period, the changes go against the potential impact of the reforms on fertility and are driven primar-

mates that the overall coverage of children under 3 by some type of childcare was 2% in 2012 (source: Eurostat, Formal child care by duration and age group).

[18]The number of nurseries declined from 1,043 in 1990 to 247 in 1993 and continued to slowly decline afterwards.

ily by external factors (for discussion, see Section 3.5.5).

3.3 Empirical strategy

3.3.1 Difference–in–differences approach

We use the 1995 and 2008 reforms of parental allowance and the difference–in–difference design to identify the effect of the duration of parental allowance on the labor market status of mothers with young children. Since the eligibility for parental allowance is universal and all women taking care of children under a given age threshold are eligible, the treatment group consists of all women with small children. In particular, we study the impact of these reforms on prime–aged (25–55) women whose youngest child was aged 2 to 7 years of age around the 1995 reform and 2 to 6 years of age around the 2008 reform.[19] We estimate the impact by the age of the youngest child in separate groups for each year of a child's age between 2 and 7 for the 1995 reform and between 2 and 6 for the 2008 reform.

We do not study the effect on mothers with children below 2 years of age, because the activity rates for these mothers are very low and stable over time (see Section 3.4) and the reforms changed eligibility for parental allowance only among women with older children. On the other hand, in contrast to most of the previous literature, we also estimate the impact on women several years after their parental leave. The oldest children in the treatment group are aged 7 years, which includes mothers of children attending the first years of primary school.[20]

The control group for the difference–in–differences strategy is fixed for all groups of treated women and consists of women whose youngest child is aged 8 to 13 years. We follow a standard approach in the literature (Naz 2004, Schone 2004, Sánchez-Mangas and Sánchez-Marcos 2008, Geyer, Haan, and Wrohlich 2014, Bergemann and Riphahn 2015) and use mothers with older children to control for the aggregate trends and business cycle effects on the labor market. While these women were likely to face similar labor market conditions as mothers with somewhat younger children, they were not affected by

[19] When studying the impact of the 2008 reform, the treatment group only includes women whose youngest child is aged 2 to 6, because mothers of 7 year old children were affected by the reform only in 2013 and the last available data are from 2012.

[20] Compulsory schooling starts at the age of 6, but the starting age is sometimes postponed till the age of 7. We do not include women with older children in the treatment group because they were not affected by the reforms in the given time frame (see below).

the reform, because they were only entitled to parental leave in the pre–reform system with their youngest child.[21]

The period of interest covers seven years of data around the 1995 reform (1993–1999)[22] and nine years of data around the 2008 reform (2004–2012). The 1995 reform affected women whose youngest child was aged 0–3 at the time of the reform (October 1, 1995) and the 2008 reform affected women whose youngest child was aged 0–2 at the time of the reform (January 1, 2008), see Section 3.2.1 for details. The earliest cohort of children whose mothers were affected by reforms were thus those born in 1992 for the 1995 reform and those born in 2006 for the 2008 reform. Therefore, the after–reform period differs for each group of treated women (defined by the age of their youngest child), as the children in the first affected cohorts age (see Table 3.1). While mothers of four–year olds in 1995 were not affected by the 1995 reform, mothers of four–year olds in 1996 and afterwards were impacted, and similarly for the other treatment groups.

Table 3.1: Definition of the after–reform period by the age of the youngest child

age of the youngest child	year					
	Panel A: The 1995 reform					
	1993–1995 Q3	1995 Q4	1996	1997	1998	1999
2	before	after	after	after	after	after
3	before	after*	after	after	after	after
4	before	before	after*	after	after	after
5	before	before	before	after*	after	after
6	before	before	before	before	after*	after
7	before	before	before	before	before	after*
	Panel B: The 2008 reform					
	2004–2007	2008	2009	2010	2011	2012
2	before	after*	after	after	after	after
3	before	before	after*	after	after	after
4	before	before	before	after*	after	after
5	before	before	before	before	after*	after
6	before	before	before	before	before	after*

Note: The Table defines the after–reform period for each group of treated women by the age of their youngest child. The asterisk denotes the first cohorts affected by each reform—women whose youngest child was born in 1992 for the 1995 reform and in 2006 for the 2008 reform (for details on the institutional background of these reforms, see Section 3.2.1).

[21] There might have been an indirect effect of the reform on their unemployment or inactivity through their possible future child. However, the likelihood that these women will have another child is not very high as only around one quarter of women have another child when their youngest child is aged above 7.

[22] The collection of the Czech LFS data started in 1993, so we cannot use earlier years.

For each group of treated women by the age of their youngest child, we estimate the following equation:

$$Y_{it} = \beta_0 + \beta_1 Treat_i + \beta_2 After_t + \beta_3 (Treat_i * After_t) + X_{it}^{'}\theta + \gamma_t + \epsilon_{it}. \tag{3.1}$$

The control group of mothers whose youngest child was aged 8–13 remains the same in all estimated equations. The outcome of interest (Y_{it}) is a binary variable that denotes a woman's labor market status, namely, non-employment (inactivity or unemployment), inactivity, and unemployment respectively. The equation includes a fixed effect for the treatment group ($Treat_i$), fixed effect for the after–reform period ($After_t$), and their interaction ($Treat_i * After_t$). The impact of parental allowance reforms is captured by β_3, which is the coefficient of the indicator variable for the treated women in the post–reform period (the period in which this group of treated women was affected by the new parental allowance legislation). We also control for the common time trend in the labor supply using fixed effects for each quarter–year combination (γ_t) and for observable characteristics (X_{it}) including quadratic polynomial of age, education dummies, dummy variables for cohabiting and married women, number of children, dummy variable for presence of elderly household members, and regional binary indicators. This equation is estimated separately for each group of treated women by the age of their youngest child, for the three labor market status outcomes (non–employment, inactivity and unemployment), and for the two reforms.

3.3.2 Data and sample selection

Our data come from the Czech Labor Force Survey (LFS), which is a quarterly survey covering about 60,000 Czech individuals. The dataset includes detailed information about household structure, economic status of the household members and their demographic characteristics, including their age. This allows us to study the labor market behavior of women by the age of their youngest child. Unfortunately, there is no information about a quarter or a month of birth for children. Therefore, we only observe age in years and do not know, for example, if a child is aged 2 years and 1 month or 2 years and 11 months. The age of the youngest child is reported in completed years—children aged 2 are children above 2 and below 3 years.[23]

[23]The survey has a rotational panel structure, which allows observation of changes in the age of the youngest child between the consecutive quarters and thus creation of a more precise measure of the

We define the economic status of women in the sample based on their self–reported status in the LFS data using the International Labour Organization (ILO) definitions. Consistently with the ILO definition, we classify an individual as unemployed if s/he does not have a job, is actively seeking a job, and is ready to start working within two weeks. However, we make one important adjustment to the ILO definition concerning maternity and parental leave. While the ILO often treats individuals on maternity and parental leave as employed, we treat them as inactive in our analysis. As we are interested in overall career interruptions after childbirth in a country with very long periods of family leave and imperfect job protection, we treat individuals on maternity and parental leave as inactive even though they may have a formal attachment to work.[24]

3.4 Changes in labor market status profiles

Before moving to the results of the regression analysis in Section 3.5, we explore the evolution of the labor market status (non–employment and unemployment) profiles of Czech mothers by the age of their youngest child over the two periods we study. The evolution of the respective knots of these profiles over time corresponds to the changes in the outcome variables that we explore in the regression analysis in Section 3.5.

3.4.1 The 1995 reform: Descriptive evidence

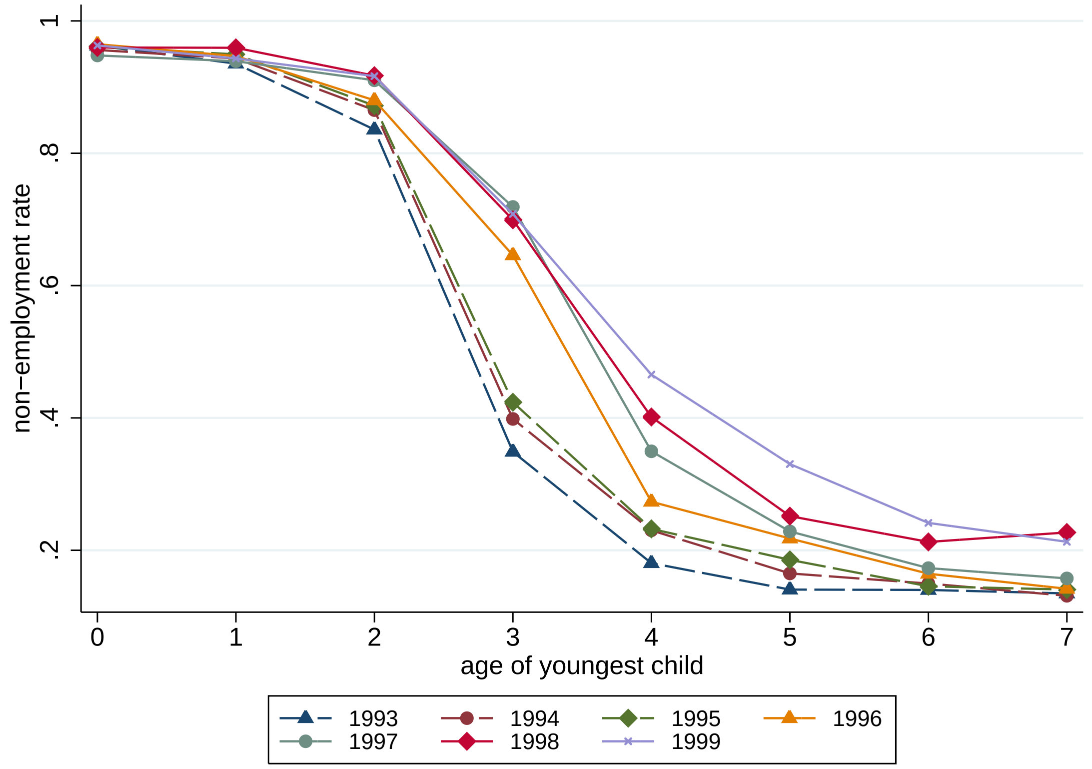
Figure 3.5 illustrates the probability that a woman does not work (the non–employment rate) by the age of woman's youngest child in years 1993–1999. The non–employment rate is almost 95% among women whose youngest child is aged 0 and 1 (till the child's first birthday, and between the first and the second birthday, respectively) and remains unaffected throughout the period. This points to a very high leave take–up rate among Czech mothers with children younger than 2 and is consistent with the fact that the first two years of family leave were not affected by the reform.[25] Throughout the studied period, the non-employment rate is somewhat lower when the child is 2, and drops more

child's age. However, this approach (used e.g. by Müllerova 2014) leads to a substantial reduction in the sample size, which would prevent us from disentangling unemployment from inactivity.

[24] The ILO definition is based on the following reasoning—if a person has a formal attachment to his/her job, but is temporarily not at work because of the maternity/parental leave, the person is still employed. While this is a reasonable assumption when the leave is short and the likelihood of a parent returning to her/his job is close to one, it still ignores the fact that while on leave, an individual does not acquire work experience but, rather, his or her human capital is likely to deteriorate.

[25] Unemployment among this group of mothers is close to zero - see Figure 3.6.

Figure 3.5: Female non–employment rate by the age of their youngest child, 1993–1999.



Note: The Figure illustrates the non–employment rate (share of inactive and unemployed in the population) for Czech women aged 25–54 by the age of their youngest child. The child's age is reported in years that the child had reached. Therefore, a child aged 3 is a child aged above 3 and below 4 years, and therefore the mother of this child is no longer covered under job protection.
Source: Czech Labour Force Survey, 1993–1999.

substantially for women whose youngest child is aged 3 and 4, as job protection expires on the child's third birthday. By the time the child is 5, only about 20% of women remain out of work, and this share remains fairly stable even for women with older children.

After the reform implemented in October 1995, which prolonged the parental allowance from 3 to 4 years, the profiles shift upwards for all groups of mothers with children aged 2 and above. The rise in non–employment is, however, most pronounced for mothers with a child aged 3. Before the reform, only around 40% of women with a child aged 3 were out of work, but this share increased to 64% in 1996 and it reached 70% in 1997 and stayed at this level for the rest of the period covered. This is not surprising, as the allowance receipt up to the child's fourth birthday was conditional on a woman's earnings and the use of childcare facilities not exceeding certain limits. The interesting fact is that the monetary motivation was stronger than the fear of losing one's job, as the job protection continued to guarantee women their jobs until a child's third birthday. Figure 3.5 also illustrates a gradual post–reform increase in the non–employment rate among women whose youngest child was aged 4 and 5. This suggests that as a result of the reform, more women stayed out of work even past their child's fourth birthday, when the prolonged parental allowance payment finished.

Figure 3.6 reveals that a large part of the post–reform increase in non–employment among women with children aged 4 and 5 was driven by an increase in their unemployment. These women were no longer covered by job protection, and when they decided to return to work they may have faced difficulty in finding a job. The unemployment–to–population rate of women whose youngest child was aged 4 gradually increased from less than 10% in the pre–reform period to 26% in 1999. The unemployment–to–population ratio also increased in the post–reform years for women with children aged 5, 6, and to some extent, 7.

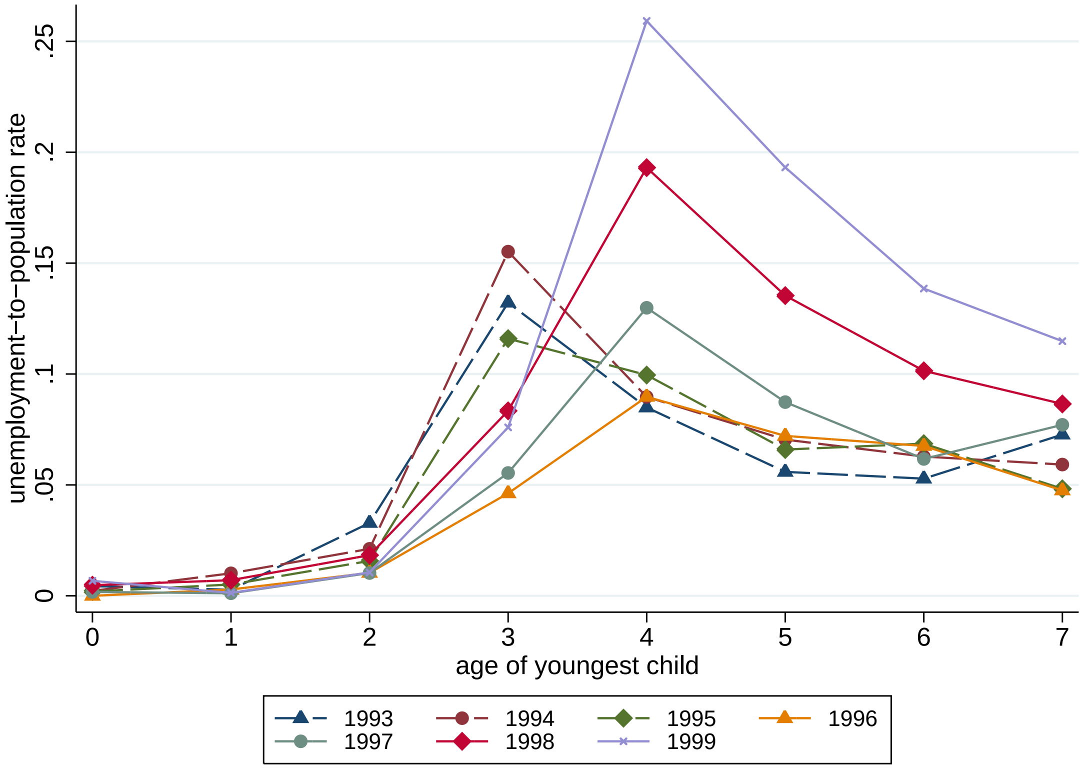
Figure 3.6 also shows that the share of unemployed women with a child aged 3 dropped in the post–reform years. This is consistent with the prolonged return to the labor market and a shift of post–leave spell of unemployment from women with 3 year old children to women with older children.[26] The gradual rise in unemployment of women with children above 4, however, might also be driven by the aggregate increase in the unemployment rate in the economy after the political and financial crises of 1997. We control for such aggregate trends with the difference–in–difference strategy in Section 3.5.

3.4.2 The 2008 reform: Descriptive evidence

We next focus on the second reform of parental allowance from January 2008, which introduced a flexible system of parental allowance that enabled some women to shorten their paid leave from four to two or three years with a higher monthly allowance amount. Figure 3.7 reports the female non–employment rate by the age of the youngest child for 2004 to 2012 and reveals that the overall shape of the non–employment profile for the pre–reform years (2004–2007) is very similar to the one shown in Figure 3.5 for 1996–1999. However, important changes in the non–employment of women with small children took place after the 2008 reform: The non–employment rate among mothers whose youngest child was aged 3 dropped from almost 70% in 2004–2007 to 60% in 2009 and then further

[26]Note that before the reform, the unemployment rate peaked among women whose youngest child was aged 3, the time when majority of women returned from their leave to the labor force. As a result of the reform, this peak has shifted till the time when the child reaches 4.

Figure 3.6: Female unemployment–to–population rate by the age of their youngest child, 1993–1999.



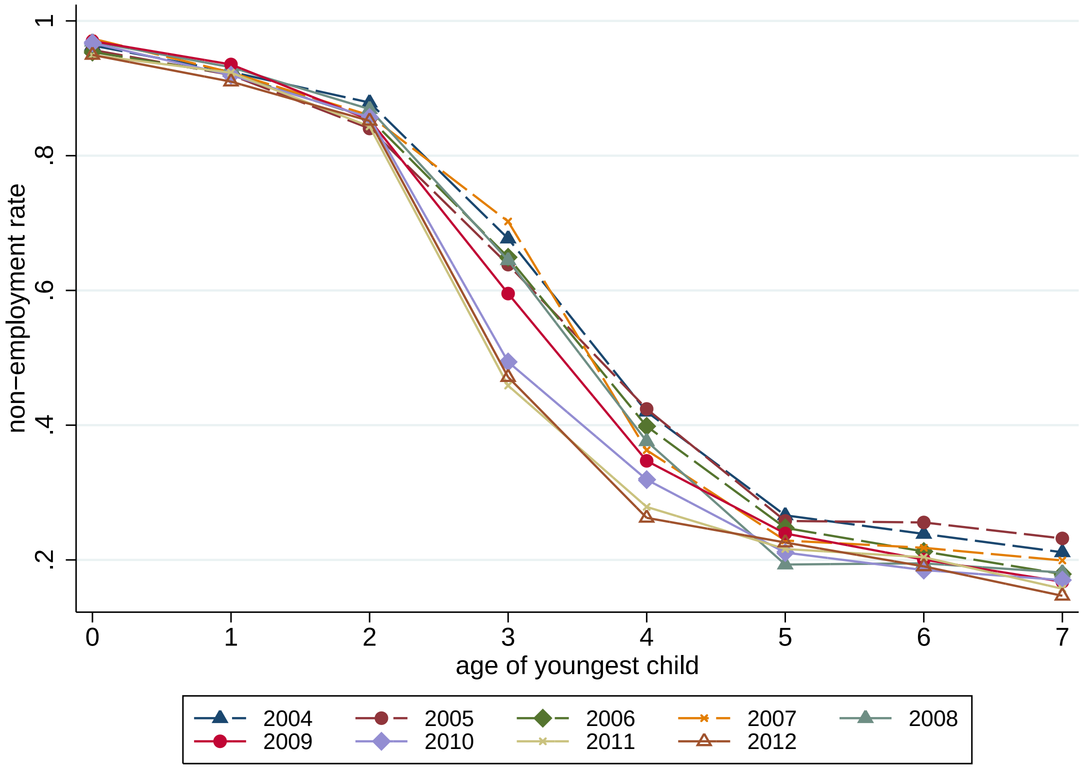
Note: The Figure illustrates the unemployment–to–population rate (share of unemployed in the population) for Czech women aged 25–54 by the age of their youngest child. The child's age is reported in years that the child had reached. Therefore, a child aged 3 is a child aged above 3 and below 4 years, and therefore the mother of this child is not under job protection anymore.
Source: Czech Labour Force Survey, 1993–1999.

to less than 50% in 2010–2012. This is consistent with a shift from a pre–reform system of four–year parental allowance to the new flexible system where some women chose a shorter track of parental allowance.[27] In addition, we also observe some decrease in the non–employment rate among women whose youngest child was aged 4 (and to a smaller extent also for children aged 5).

At the same time, Figure 3.8 shows that the unemployment–to–population rate for these women did not change much in the post–reform period (it already decline in 2006), so it seems that the decrease in the non–employment rate of these women was driven by a drop in inactivity. The unemployment–to–population rate increased substantially only for women whose youngest child was aged 3 in the post–reform period (Figure 3.8).

[27] The delay in the effects of the reform was caused by its design. The first women who took advantage of the flexible system returned to the labor market only in 2009 (for details, see Section 3.2.1).

Figure 3.7: Female non–employment rate by the age of their youngest child, 2004–2012.

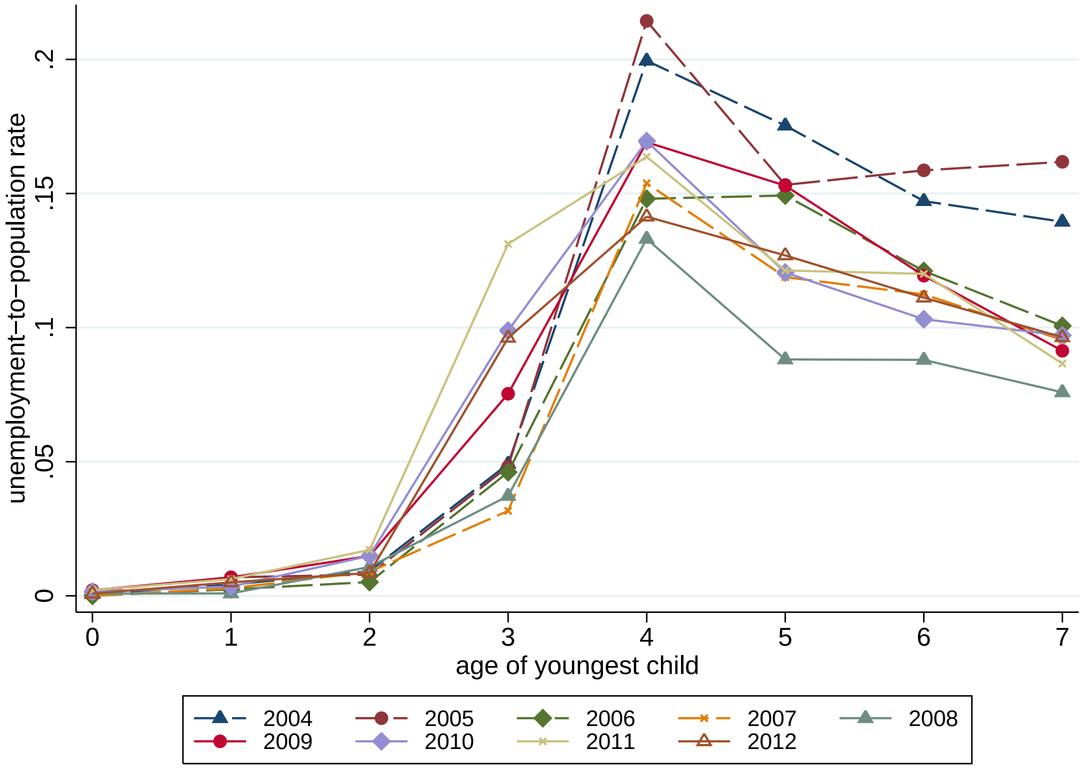


Note: The Figure illustrates the non–employment rate (share of inactive and unemployed in the population) for Czech women aged 25–54 by the age of their youngest child. The child's age is reported in years that the child had reached. Therefore, a child aged 3 is a child aged above 3 and below 4 years, and therefore the mother of this child is not under job protection anymore.
Source: Czech Labour Force Survey, 2004–2012.

However, the development of unemployment in this period might be confounded by the economic crisis that hit the Czech labor market in 2009. We deal with this problem by filtering out the effect of aggregate labor market conditions with control groups in the next section.

The 2008 reform allowed some women to choose the two–year track of parental allowance. However, we see no evidence in the data that women were more likely to return to work after two years on leave in the post–reform period (see Figures 3.7 and 3.8). This is consistent with the statistics of the Czech Ministry of Labour and Social Affairs, which report that only 6% of women chose the two–year track in 2011. The main reason Czech women tend to stay on leave for at least three years is that, as noted earlier, public childcare facilities for children below 3 are virtually nonexistent in the Czech Republic and private ones are quite expensive (see Section 3.2.3).

Figure 3.8: Female unemployment–to–population rate by the age of their youngest child, 2004–2012.



Note: The Figure illustrates the unemployment–to–population rate (share of unemployed in the population) for Czech women aged 25–54 by the age of their youngest child. The child's age is reported in years that the child had reached. Therefore, a child aged 3 is a child aged above 3 and below 4 years, and therefore the mother of this child is not under job protection anymore.
Source: Czech Labour Force Survey, 2004–2012.

3.5 Estimation results

3.5.1 Baseline difference–in–differences results

Tables 3.2 and 3.3 present the main set of results. In our baseline specification, we estimate the effect of the 1995 and 2008 reforms on a mother's labor market status after childbirth by the age of the youngest child up to the age of 6/7, using mothers with older children (8–13) to control for overall changes in the labor market via the difference–in–difference design.[28] Our findings confirm the patterns already discernible from the labor market status profiles presented in Section 3.4, when cleared from aggregate trends and business cycle effects.

[28]Appendix Tables 3.8 and 3.9 show sample summary statistics for the treatment and control groups, before and after the 1995 and 2008 reforms.

The 1995 reform, which extended the duration of parental allowance from three to four years (but kept the job protection at three years), increased non–employment among all mothers with children between 3 and 6 years of age. The results show that after the reform, mothers were 25.5 p.p. more likely to be at home when their child was 3 than before the reform (see Panel A in Table 3.2), in spite of the fact that the job protection period elapsed at the child's third birthday.[29] This either implies greater importance of the monetary aspect of the parental leave over job security or questions the actual strength of the job protection.[30]

The rise in non–employment of mothers induced by the reform diminished with the youngest child's age but was still fairly high at the age of 4 and 5—at 11.3 and 4.2 p.p., respectively. We observe no effect on women whose children had already turned 7 (when virtually all children are already at school) suggesting that the impact of the reform on non–employment gradually fades away.

Decomposition of the effect of the 1995 reform on the non–employment rate into the effect on inactivity and on unemployment (Panels B and C of Table 3.2) reveals that the impact of the parental allowance extension on overall career breaks after childbirth operated through the following channels: The reform substantially raised the inactivity of all women with children between 2 and 4, while shifting the after–leave unemployment from ages 2–3 towards 4–6. The impact of the reform on women's probability of staying at home for at least one more year was enormous: due to the reform the probability of a woman with a child between 3 and 4 being inactive increased by as much as 35.6 p.p. The effect on mothers with children aged 4 is a combination of prolonged inactivity (6.4 p.p.) and increased unemployment (4.9 p.p.). For mothers of older children (aged 5 and 6), the increase in non–employment is entirely driven by a rise in unemployment. The share of women who were inactive or unemployed with a child of the age of 7 or more remained unaffected by the reform.

As expected, the 2008 reform that offered some women (but not all) an option to receive the same overall parental leave allowance over a shorter period of two or three

[29] Our non–employment results are thus strikingly similar to the estimates from Müllerova (2014) who finds—using a regression discontinuity approach—that the employment probability of mothers with 3 year old children fell by 23 p.p.

[30] While some women may not be entitled to the job protection, given the rather weak eligibility requirements, the jobs of the majority of women are under the protection. As we discuss in Section 3.2.1, however, several sociological studies pointed out that the job protection may be weak in labor markets in transition (Kantorova 2004, Fodor 2005) and there is evidence of job protection avoidance by Czech employers even nowadays (Kucharova 2006).

Table 3.2: Results of the difference–in–differences estimation: 1995 reform

	Treatment group: women whose youngest child is:					
	aged 2	*aged 3*	*aged 4*	*aged 5*	*aged 6*	*aged 7*
			Panel A: impact on non–employment			
Treat*After	0.011	0.255***	0.113***	0.042***	0.021***	0.007
	(0.007)	(0.016)	(0.014)	(0.007)	(0.006)	(0.007)
R-squared	0.436	0.22	0.09	0.065	0.056	0.057
Observations	83548	83702	84102	84290	84475	84403
			Panel B: impact on inactivity			
Treat*After	0.047***	0.356***	0.064***	0.002	-0.001	-0.002
	(0.009)	(0.016)	(0.006)	(0.006)	(0.003)	(0.005)
R-squared	0.535	0.241	0.048	0.034	0.031	0.034
Observations	83548	83702	84102	84290	84475	84403
			Panel C: impact on unemployment			
Treat*After	-0.036***	-0.101***	0.049***	0.040***	0.022***	0.009*
	(0.005)	(0.006)	(0.012)	(0.007)	(0.005)	(0.005)
R-squared	0.031	0.035	0.051	0.042	0.037	0.035
Observations	83548	83702	84102	84290	84475	84403

Note: The treatment groups consist of prime–aged women (aged 25–55), whose youngest child is aged 2–7. The control group consists of prime–aged women whose youngest child is aged 8–13. For each treatment group, a separate regression is estimated by the age of the youngest child, but the control group is fixed in all regressions. All regressions include dummies for the treatment group and after period, quarter–year dummies, and other control variables. The after period is defined differently in each regression for both treatment and control groups according to the quarter in which children of women who were eligible for the post–reform parental allowance with this child reached that particular age. Standard errors (in parentheses) are clustered at the group–year level (* $p<0.10$, ** $p<0.05$, *** $p<0.01$). Source: Czech LFS (1993–1999), own calculations.

years, had an opposite but smaller effect. Panel A in Table 3.3 shows that an option to reduce the duration of paid parental leave led to a decrease in non–employment by 14 and 7.6 p.p., of mothers of 3 and 4 year olds, respectively. It raised, however, the probability of non–employment of mothers of 2 year olds by 3.6 p.p. These results can again be partly explained by a “trade–of” between inactivity and unemployment, as in the case of the 1995 reform. Panels B and C of Table 3.3 reveal that the 2008 reform reduced the probability of inactivity of a woman with a 3 and 4 year old child by as much as 21.4 and 7.2 p.p. respectively but had no impact on inactivity of other mothers. The reduction in inactivity, however, was partly compensated by a rise in the likelihood of unemployment at earlier stages, when the child is 2 and 3 (3.0 and 7.2 p.p. increase).

Table 3.3: Results of the difference–in–differences estimation: 2008 reform

	Treatment group: women whose youngest child is:				
	aged 2	*aged 3*	*aged 4*	*aged 5*	*aged 6*
	Panel A: impact on non–employment				
Treat*After	0.036***	-0.142***	-0.076***	0.005	-0.013
	(0.006)	(0.014)	(0.008)	(0.006)	(0.008)
R-squared	0.448	0.241	0.126	0.099	0.099
Observations	78768	76727	75225	74251	74001
	Panel B: impact on inactivity				
Treat*After	0.006	-0.214***	-0.072***	0.009*	-0.002
	(0.005)	(0.018)	(0.008)	(0.005)	(0.005)
R-squared	0.586	0.288	0.059	0.034	0.034
Observations	78768	76727	75225	74251	74001
	Panel C: impact on unemployment				
Treat*After	0.030***	0.072***	-0.004	-0.004	-0.011*
	(0.004)	(0.007)	(0.007)	(0.004)	(0.006)
R-squared	0.066	0.056	0.073	0.071	0.072
Observations	78768	76727	75225	74251	74001

Note: The treatment groups consist of prime–aged women (aged 25–55), whose youngest child is aged 2–6. The control group consists of prime–aged women whose youngest child is aged 8–13. For each treatment group, a separate regression is estimated by the age of the youngest child, but the control group is fixed in all regressions. All regressions include dummies for the treatment group and after period, quarter–year dummies, and other control variables. The after period is defined differently in each regression for both treatment and control groups according to the quarter in which children of women who were eligible for the post–reform parental allowance with this child reached that particular age. Standard errors (in parentheses) are clustered at the group–year level (* $p<0.10$, ** $p<0.05$, *** $p<0.01$). Source: Czech LFS (2004–2012), own calculations.

In other words, the flexible parental leave allowance schedule introduced in 2008 induced some mothers to take shorter family leave, and therefore also to experience the potential post–leave unemployment earlier (when their children are younger) than prior to the reform. Yet, the reform did not reduce the probability of unemployment among mothers with children older than 4. We attribute this finding to the fact that it is the low–income mothers who have the greatest difficulty in finding a job after the leave, but who are at the same time most likely to take the three–year leave or to continue with the four–year leave, and who thus might have been unaffected by the reform.

Under the assumption that women do not return to parental leave once they have entered the labor force after childbirth, we can calculate the impact of the two reforms on the average leave duration using the pre–reform inactivity rates and our baseline specification estimates.[31] We find that while the 1995 reform extended the average leave duration from 2.8 to 3.3 years, the 2008 reform shortened it again from 3.2 to 2.9 years.

In sum, the 1995 reform extended the family leave duration and led to a rise in the probability of female unemployment at higher ages of children, thus substantially prolonging the overall after–birth career breaks of the majority of mothers in the Czech Republic to more than three years. The 2008 reform, on the other hand, achieved a reversal of the impact of the 1995 reform, but only to some extent.

The impact of the two reforms on the post–birth labor market status of women, however, disappears by the time the youngest child turns 7 (in case of the 1995 reform) and 5/6 (in case of the 2008 reform). Future research is needed to explore the potential consequences of these medium run effects of the parental leave reforms on earnings and the quality of jobs women hold after childbirth, and whether these effects persist over time. Unfortunately, the data currently available have neither information about income, nor the panel dimension that would allow us to compare pre–birth and post–birth jobs.

3.5.2 Supplementary evidence on leave to unemployment transitions

Our baseline results suggest that the period of a post–birth leave is often followed by a spell of unemployment. This subsection studies the channels through which mothers with small children become unemployed. The job protection mechanism should ensure that women can return to their jobs after the period of family leave without the risk of becoming unemployed. However, some mothers may not be entitled to the job protection and some may extend their family leave beyond the protected period. Moreover, there is some evidence questioning the strength and enforceability of the job protection. Job protection is weak in the labor markets in transition (as in 1990s in the Czech Republic, see e.g. Kantorova 2004, Fodor 2005) or at times of economic turmoil, because pre–birth jobs previously held by mothers could have been eliminated during the several–year long

[31] We calculate the leave duration conditional on it lasting between 1 and 7 years. This seems to be a reasonable assumption, as 96% of Czech women with a child below 1 year of age are inactive and only around 7% of mothers are inactive by the time their child reaches 7 years of age. In particular, we calculate the average leave duration as $1 * (IR_0 - IR_1) + 2 * (IR_1 - IR_2) + ... + 7 * (IR_6 - IR_7)$, where IR_a is the inactivity rate of women whose youngest child is aged a. The inactivity rates are taken from the data for the before–reform period, and calculated using the baseline specification estimates reported in Tables 3.2 and 3.3 for the after–reform period.

family leave. The need to search for a job immediately following the family leave then results in a direct transition from leave to unemployment.[32]

Alternatively, the unemployment spell may not follow immediately but within a very short period of time after the family leave. The negative impact of a lengthy career break on human capital and productivity make women who return to their jobs after family leave more likely to become unemployed soon afterwards.[33]

In what follows, we look at the share of unemployed mothers who entered unemployment directly from leave (see Figure 3.9) as opposed to those who entered unemployment after a spell of employment. Clearly, the majority of women with children below 4 enter unemployment directly from leave suggesting that they either have no job to return to or the employer somehow avoids provision of job protection. This share decreases with a child's age, meaning that women with older children most often become unemployed after a (short) spell of employment.

The 1995 reform substantially increased the probability that a women enters unemployment directly from leave for all mothers with children between 2 and 7 years of age, but especially for those with a 4 year old child, for whom this probability more than doubled. This is in line with our expectations, as the reform increased the share of mothers who extended their family leave beyond the job–protected period. The 2008 reform, on the other hand, decreased the share of unemployed women who enter unemployment directly from leave, but only to a smaller extent and only for women with children above 4 years of age.

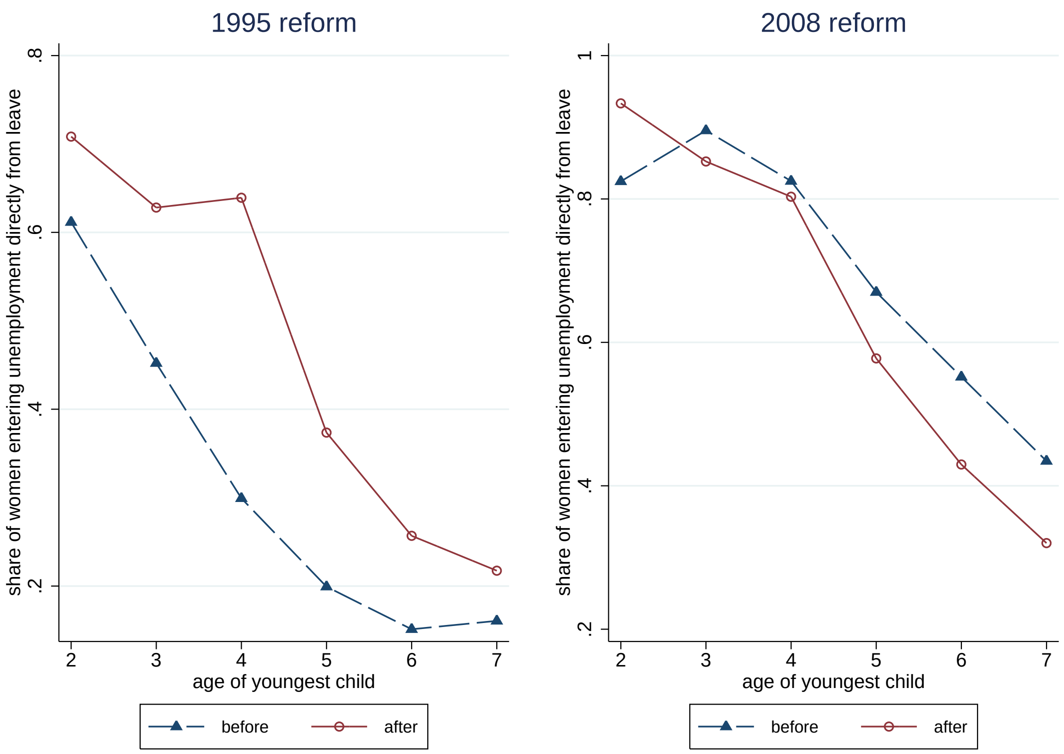
3.5.3 Difference–in–differences results by education

The flexibility of the parental leave allowance schedule introduced in 2008 depends on income. Moreover, both reforms may induce unequal responses from women with different earnings potential. Lalive and Zweimüller (2009) note that monetary aspects of the parental leave are likely to be more important for the labor market status of lower–income women (given higher replacement rates), whereas job protection is a greater concern for

[32] This so called mechanical impact of family leave on unemployment in the absence of job protection was first pointed out by Johnson (1983).

[33] Over one third of Czech women who do not continue with their pre–birth employment point to their employer's adverse attitude towards their return to work as the main reason for leaving the job (Kucharova 2006). International evidence shows that weak attachment to the pre–birth job after the family leave, potentially driven by employers' attempts to go around the job protection regulations, is likely to be widespread. E.g. Schönberg and Ludsteck (2014) report that 7% of women who return to work at the end of the statutory leave become unemployed 1 or 2 months afterwards.

Figure 3.9: Share of unemployed women entering unemployment directly from leave



Note: The Figure illustrates the share of unemployed women who report being on maternity or parental leave directly before they entered unemployment. A child's age is reported in years that the child had reached. Therefore, a child aged 3 is a child aged above 3 and below 4 years, and therefore the mother of this child is not under job protection anymore.
Source: Czech Labour Force Survey, 1993–1999 and 2004–2012.

higher–income women (for example because of the career concerns). We would therefore expect the 1995 reform to have a greater impact on low–income women. The expectations of the impact of the 2008 reform are more complex, as only high–income women (or women with a high–income husband) were eligible for the option of the shortest duration of parental allowance with the highest monthly payment.

In what follows, we use the best proxy for income we have in our data and explore the potential heterogeneity in women's responses to the two reforms by estimating our baseline specification separately by two levels of education: low and high.[34] Before we discuss the estimation results, we note that there are remarkable similarities between low

[34]Unfortunately, this is the best proxy for income we have in the data. 'Low–educated' corresponds to the ISCED 3 level with apprenticeship certificate, but without school leaving examination, or lower level of education (corresponding to less than A levels in the UK, without a Baccalaureate in France or without a high–school diploma in the US). Finer classification using ISCED renders a too small sample size for some of the groups.

and high educated women in terms of parental leave take–up, as measured by the share of inactive mothers of children between 2 and 7 years old (see Table 3.4). The key differences in the non–employment rate between these two groups are driven predominantly by a different likelihood of unemployment, which is twice as high (three times as high) among the low educated when compared to the high educated prior to the 1995 (2008) reform.

Table 3.4: Summary statistics by woman's education

	1995 reform				2008 reform			
	High education		Low education		High education		Low education	
	Before	After	Before	After	Before	After	Before	After
Non–employed	0.269	0.345	0.353	0.486	0.393	0.371	0.534	0.492
Inactive	0.215	0.284	0.258	0.371	0.334	0.315	0.377	0.346
Unemployed	0.053	0.061	0.095	0.115	0.059	0.056	0.157	0.146
Observations	14475	22268	13787	21149	18497	25350	15122	16127

Notes: The sample includes all treated women, i.e. women whose youngest child is aged 2-7 years. High education corresponds to ISCED 3 level with school leaving examination or more, while low education is defined as ISCED 3 level with apprenticeship certificate (but without school leaving examination) or less. The before period for the 1995 reform is defined as 1993 Q1–1995 Q3 and the after period as 1995 Q4–1998 Q4. The before period for the 2008 reform covers years 2004–2007 and the after period years 2008–2012. Source: Czech LFS (1993–2012), own calculations.

The results of difference–in–differences estimation, when the reform effect is interacted with the level of educational, confirm that the one year extension of parental allowance in the 1995 reform affected the probability of being inactive with a child between 3 and 4 among low educated women more (41 p.p. increase) than among the high educated women (30 p.p. increase), suggesting that the benefit receipt mattered more for the low–educated (see Panel B of Table 3.5).[35] Similarly, the inactivity among the mothers of 3 year olds dropped somewhat more for the low educated women (23 p.p.) than for the high educated (19 p.p.) in response to the 2008 reform (see Panel B of Table 3.6).

The impact on unemployment (shifting the risk of unemployment from earlier stages for mothers with children below 4 years to later stages after the 1995 reform and increasing unemployment of mothers with 2 and 3 year old children in 2008) was somewhat greater for the low educated (see Panel C in Tables 3.5 and 3.6), who in general face a higher risk of unemployment. The smaller differences in the impact of the 2008 reform across the two education groups when compared to the 1995 reform are likely to be driven by the fact that some of the low–income mothers did not have the option to choose a shorter

[35]The reform, however, also induced high–educated mothers with 2 year old children to be more inactive than the low educated.

Table 3.5: Difference–in–differences estimation by education: 1995 reform

	Treatment group: women with the youngest child					
	aged 2	*aged 3*	*aged 4*	*aged 5*	*aged 6*	*aged 7*
			Panel A: impact on non–employment			
Treat*After	-0.016*	0.285***	0.097***	0.032**	0.048***	0.024**
	(0.008)	(0.015)	(0.019)	(0.013)	(0.011)	(0.012)
Treat*After*HighEduc	0.052***	-0.058***	0.035*	0.02	-0.052***	-0.035**
	(0.015)	(0.014)	(0.018)	(0.017)	(0.017)	(0.018)
R-squared	0.44	0.226	0.096	0.07	0.062	0.064
Observations	83544	83698	84098	84286	84470	84399
			Panel B: impact on inactivity			
Treat*After	0.028***	0.411***	0.058***	-0.006	0.004	-0.007
	(0.008)	(0.017)	(0.010)	(0.011)	(0.005)	(0.009)
Treat*After*HighEduc	0.037**	-0.108***	0.014	0.015	-0.01	0.01
	(0.015)	(0.016)	(0.012)	(0.013)	(0.009)	(0.011)
R-squared	0.537	0.247	0.051	0.037	0.033	0.037
Observations	83544	83698	84098	84286	84470	84399
			Panel C: impact on unemployment			
Treat*After	-0.043***	-0.126***	0.039**	0.038***	0.044***	0.032***
	(0.006)	(0.006)	(0.016)	(0.011)	(0.010)	(0.009)
Treat*After*HighEduc	0.015**	0.050***	0.021	0.004	-0.042***	-0.045***
	(0.007)	(0.012)	(0.014)	(0.013)	(0.014)	(0.013)
R-squared	0.034	0.037	0.053	0.045	0.041	0.039
Observations	83544	83698	84098	84286	84470	84399

Note: The treatment groups consist of prime–aged women (aged 25–55), whose youngest child is aged 2–7. The control group consists of prime–aged women whose youngest child is aged 8–13. For each treatment group, a separate regression is estimated by the age of the youngest child, but the control group is fixed in all regressions. All regressions include dummies for the treatment group and after period, quarter–year dummies, and other control variables. The after period is defined differently in each regression for both treatment and control groups according to the quarter in which children of women who were eligible for the post–reform parental allowance with this child reached that particular age. Standard errors (in parentheses) are clustered at the group–year level (* p<0.10, ** p<0.05, *** p<0.01). Source: Czech LFS (1993–1999), own calculations.

paid parental leave duration and were therefore unaffected by the 2008 reform.

3.5.4 Robustness analysis

The robustness analysis presented in this section shows results of the difference–in–differences estimation with an alternative definition of the control group. The choice of the control group here was motivated by the fact that the control group in the baseline specification is defined by the age of the youngest child and thus includes different cohorts

Table 3.6: Difference–in–differences estimation by education: 2008 reform

	Treatment group: women with the youngest child				
	aged 2	*aged 3*	*aged 4*	*aged 5*	*aged 6*
	Panel A: impact on non–employment				
Treat*After	0.046***	-0.149***	-0.059***	0.026**	-0.009
	(0.009)	(0.013)	(0.016)	(0.011)	(0.009)
Treat*After*HighEduc	-0.025**	0.016	-0.018	-0.036***	-0.005
	(0.011)	(0.017)	(0.019)	(0.013)	(0.018)
R-squared	0.453	0.25	0.137	0.108	0.108
Observations	78767	76726	75223	74249	74000
	Panel B: impact on inactivity				
Treat*After	0.013*	-0.231***	-0.057***	0.018*	0
	(0.007)	(0.018)	(0.009)	(0.010)	(0.008)
Treat*After*HighEduc	-0.01	0.037**	-0.023**	-0.017	-0.005
	(0.010)	(0.016)	(0.009)	(0.011)	(0.012)
R-squared	0.587	0.295	0.062	0.036	0.036
Observations	78767	76726	75223	74249	74000
	Panel C: impact on unemployment				
Treat*After	0.033***	0.082***	-0.003	0.007	-0.009
	(0.006)	(0.011)	(0.016)	(0.009)	(0.007)
Treat*After*HighEduc	-0.015**	-0.022*	0.006	-0.018	0
	(0.007)	(0.011)	(0.020)	(0.012)	(0.010)
R-squared	0.076	0.062	0.081	0.077	0.079
Observations	78767	76726	75223	74249	74000

Note: The treatment groups consist of prime–aged women (aged 25–55), whose youngest child is aged 2–6. The control group consists of prime–aged women whose youngest child is aged 8–13. For each treatment group, a separate regression is estimated by the age of the youngest child, but the control group is fixed in all regressions. All regressions include dummies for the treatment group and after period, quarter–year dummies, and other control variables. The after period is defined differently in each regression for both treatment and control groups according to the quarter in which children of women who were eligible for the post–reform parental allowance with this child reached that particular age. Standard errors (in parentheses) are clustered at the group–year level (* p<0.10, ** p<0.05, *** p<0.01). Source: Czech LFS (2004–2012), own calculations.

of women at different points in time. The control group at the beginning of the period of interest might thus include women who were affected by different policies than those who are in the control group at the end of the period. Therefore, if previous reforms of parental allowance had an influence on the long–term labor market outcomes of mothers in the control group, we have to move the control group to earlier cohorts, where there were no reforms of parental allowance. The alternative control group used in the robustness analysis is thus defined to meet this condition—it includes women whose youngest child was aged 13–23 for the 1995 reform, and women whose youngest child was aged 18–25 for the 2008 reform.[36]

As suggested by Appendix Tables 3.7 and 3.10, the results are largely confirmed when we use this alternative control group of women with older children. The main differences are as follows: The robustness check for the 1995 reform implies, first, that the extension of the parental allowance by one year increased inactivity and unemployment among mothers with a child older than 4 more and reduced unemployment among mothers with younger children less than suggested by the baseline specification, and, second, that the impact of the reform on the overall career break of mothers after childbirth may extend beyond the medium run, in the form of a 2.4 p.p. increase in the unemployment of women with children aged 7 (see Table 3.7). The robustness check for the 2008 reform, on the other hand, suggests an even greater reduction in the probability of unemployment at later child age stages (including the age of 5 and 6) than estimated earlier (see Table 3.10).

Our baseline specifications therefore provide more conservative estimates of the main effects of the two reforms than implied by the sensitivity analysis. We conjecture that the relatively small dissimilarities are mainly driven by the fact that women with older children (who constitute the alternative control group) differ more substantially from both our treatment as well as our baseline control group in terms of lower sensitivity of their labor market outcomes to changes in the economy.[37]

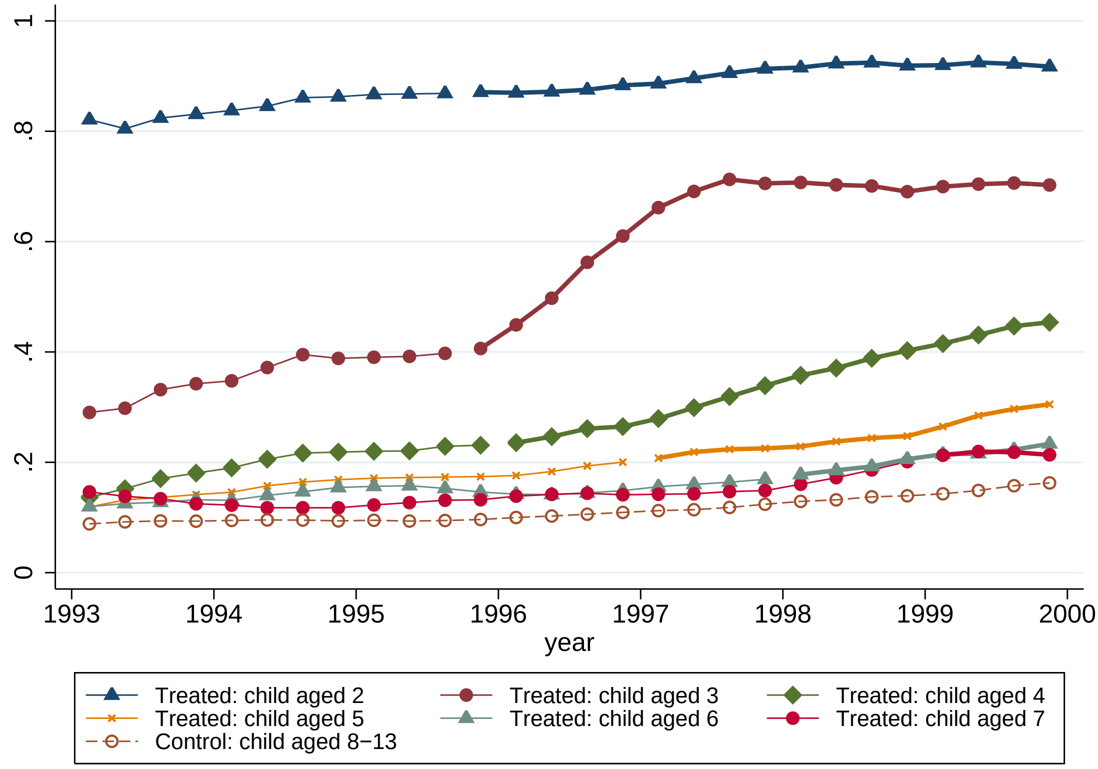
3.5.5 Identification assumptions

Our empirical strategy is based on the assumption of common trends of the treatment and control groups, i.e. we assume that trends in labor market status (non–employment and unemployment) of the treatment group of women whose youngest child is aged 2–7 and of the control group with youngest children aged 8–13 would have been the same absent of any treatment. To provide some evidence regarding the validity of this assumption, we plot the evolution of the non–employment and unemployment–to–population rates of all the treatment groups and the control group over the two studied periods in Figures

[36]Women in the alternative control group for the 1995 reform have a youngest child born between 1970 and 1985—the period during which no major reform of parental leave took place. Women in the alternative control group for the 2008 reform gave birth to their youngest child between 1979 and 1992, i.e. in the pre–1995 parental allowance system.

[37]The evolution of the labor market status of women in the two alternative control groups, however, is basically parallel over the two periods that we study. Figures are not included due to space limitations but are available from the authors upon request.

Figure 3.10: Non–employment rate of women by the age of youngest child, 1993–1999.



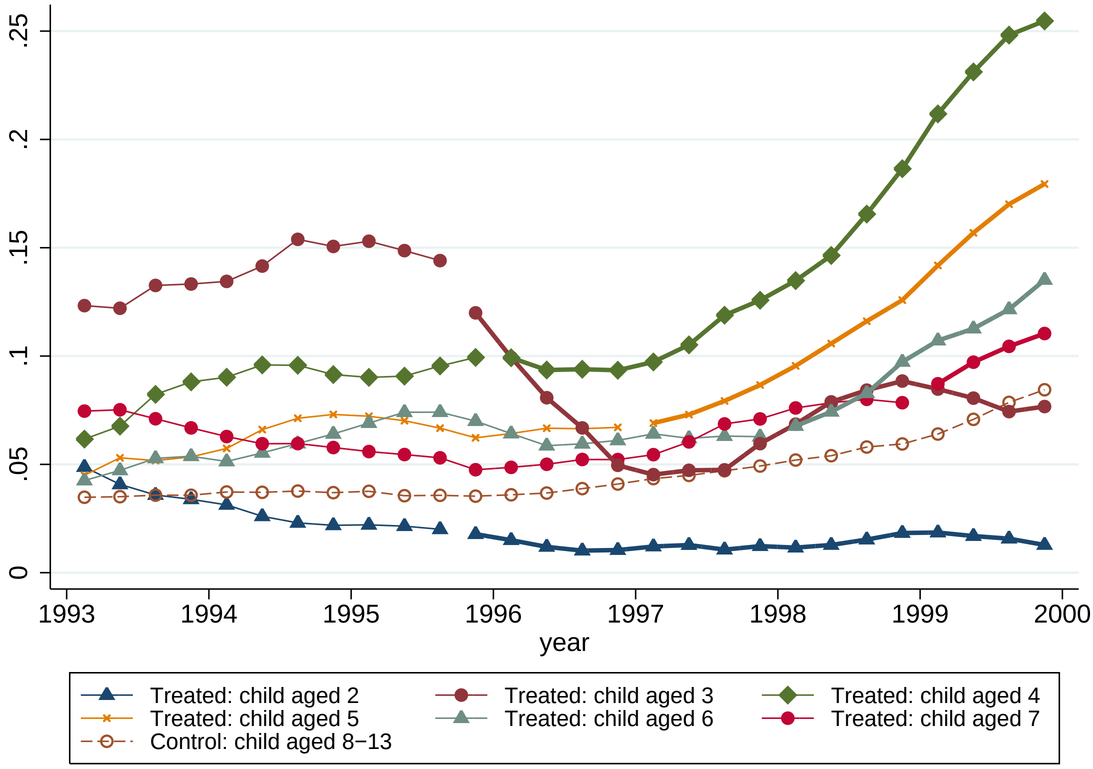
Note: The figure depicts the non–employment rate (share of inactive and unemployment in the population) for women in the treatment and control groups by the age of their youngest child. The after reform period is denoted by a thick line for each group of treated women. The time series were seasonally adjusted using MA(4) smoothing. Source: Czech LFS data (1993–1999).

3.10–3.13. The after period is denoted by a thick line for each group of treated women by the age of their youngest child.[38] To abstract from seasonality and allow a better focus on aggregate trends, all time series presented in this section were seasonally adjusted using standard MA(4) smoothing.

The evolution over the first period, covering the 1995 reform, is presented in Figures 3.10 and 3.11. The non–employment rate of women whose youngest child is aged 2 seems to increase slightly over the pre–treament period (Figure 3.10), which is consistent with the evolution of the control group (women whose youngest child is aged 8–13). Women with children aged 3 experienced a more pronounced increase in their non–employment rate in the early pre–treatment period, but this stabilized in the fourth quarter of 1994

[38]The after–reform period differs for each group of treated women defined by the age of their youngest child according to the quarter in which children of women who were eligible for the post–reform parental allowance with this child reached that particular age (for details on after period definition, see Section 3.3.1 and Table 3.1).

Figure 3.11: Unemployment–to–population rate of women by the age of youngest child, 1993–1999.

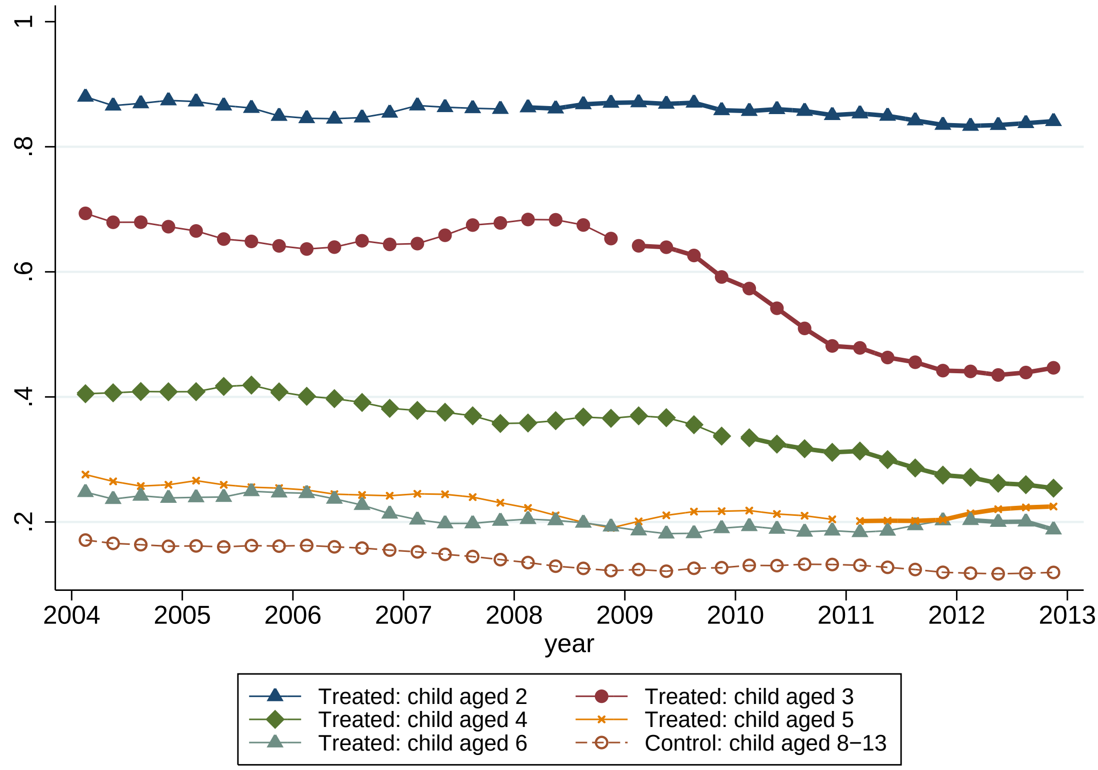


Note: The figure depicts the unemployment–to–population rate (share of unemployment in the population) for women in the treatment and control groups by the age of their youngest child. The after reform period is denoted by a thick line for each group of treated women. The time series were seasonally adjusted using MA(4) smoothing. Source: Czech LFS data (1993–1999).

and did not change much in the rest of the pre–reform period. The evolution of the non–employment rate seems to follow the control group very closely for all remaining groups of treated women (with children aged 4, 5, 6, and 7) in the pre–treatment period, i.e. the period for which the line is thin.

Figure 3.11 provides some evidence for the validity of the common trend assumption for the unemployment–to–population rate. The pre–treatment trend for women with children aged 2 seems to differ from the control group, as these women experienced a gradual decrease in their unemployment probability, which was a likely consequence of an increasing share of inactive women in this group (see Figure 3.10). For women with slightly older children (aged 3 and 4), unemployment increased somewhat in 1993 and 1994, while the control group's unemployment was relatively stable over the whole pre–treatment period. The unemployment–to–population rate for mothers with children aged 5, 6, and 7 follow the overall evolution of the control group quite well, apart from an increase in the unemployment rate of women with children aged 5 in 1994 and women with children aged 6 in 1995. However, their unemployment decreased again and stabilized afterwards, so that it followed relatively closely the evolution of the control group for the rest of the pre–treatment period (up until 1997 and 1998, respectively). Therefore, the validity of the common trend assumption might be questioned for some groups of treated women, but the divergences in trends are quite small compared to the observed impacts of the reform.

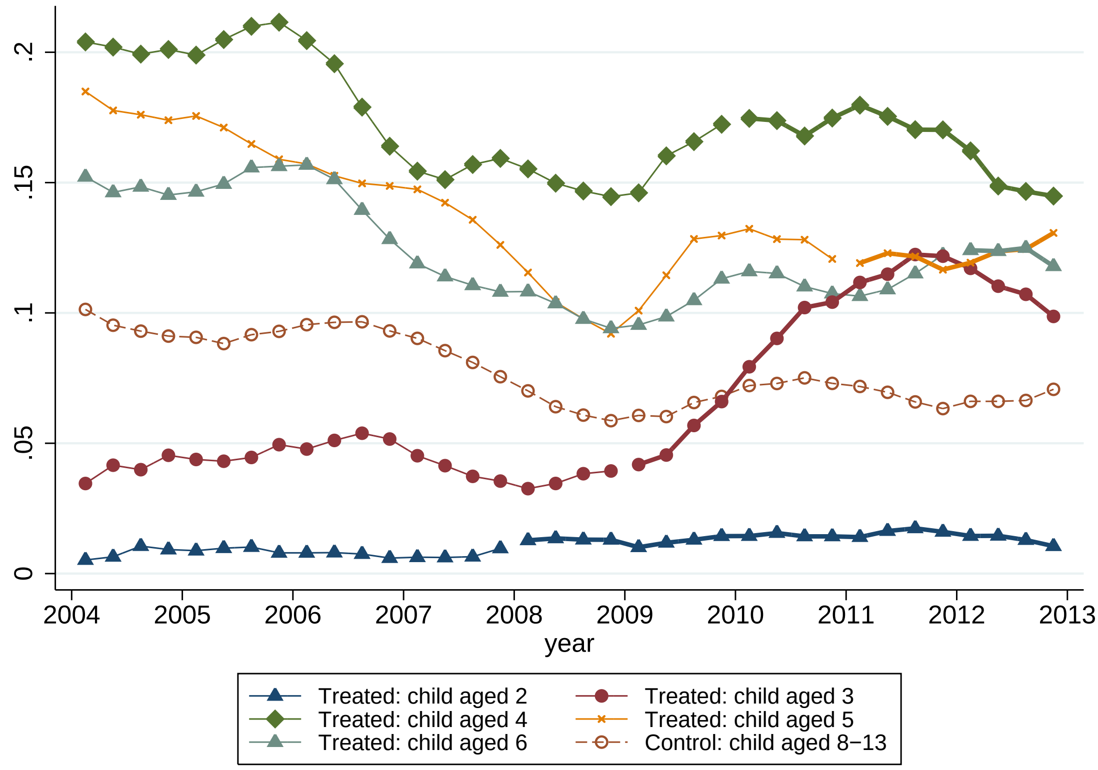
Figure 3.12: Non–employment rate of women by the age of youngest child, 2004–2012.



Note: The figure depicts the non–employment rate (share of inactive and unemployed in the population) for women in the treatment and control groups by the age of their youngest child. The after reform period is denoted by a thick line for each group of treated women. The time series were seasonally adjusted using MA(4) smoothing. Source: Czech LFS data (2004–2012).

The common trend assumption for the 2008 reform is investigated in Figures 3.12 and 3.13. The non–employment trends for women with children aged 2–6 are quite stable or show some mild decrease in the pre–treatment period, but are quite similar to the control group of women with children aged 8–13 (see Figure 3.12). The share of unemployed women with children aged 2 is very low (below 2%) and relatively stable over the whole pre–reform period (Figure 3.13). Women with children aged 3, 4, and 6 experienced some increase in unemployment probability in the second half of 2005 or in 2006 and a subsequent decline, which was again consistent with the evolution of the control group. Therefore, the common trend assumption seems to be quite reasonable for both the non–employment and unemployment–to–population rates around the 2008 reform.

Figure 3.13: Unemployment–to–population rate of women by the age of youngest child, 2004–2012.



Note: The figure depicts the unemployment–to–population rate (share of unemployment in the population) for women in the treatment and control groups by the age of their youngest child. The after reform period is denoted by a thick line for each group of treated women. The time series were seasonally adjusted using MA(4) smoothing. Source: Czech LFS data (2004–2012).

The empirical strategy further requires that there were no significant composition changes in the treatment and control groups. This assumption could be violated if fertility decisions of Czech couples were significantly influenced by the reforms of the parental allowance, and the fertility changes then affect the composition of the treatment and control groups.

Fertility changes in the 1990s and early 2000s were described in Section 3.2.3 and seem to be unrelated to the reforms of parental allowance. The steep decline in fertility rates in the 1990s started long before the 1995 reform and was part of a general trend that decreased fertility in all post–communist countries (Sobotka 2003). Moreover, this trend goes against the expected impact of the parental allowance reform on fertility—the more generous parental allowance should (if anything) promote and incentivize parenthood. Fertility changes in the 2000s also took place before the 2008 reform and seem to be unrelated to it. However, there might still be a concern that the population waves might have changed the composition of women affected by the reform. To address this potential issue, Appendix Table 3.11 reports characteristics of Czech mothers with a new–born child (child aged 0) from 1993 to 2012. Over time, mothers of new–born children have become slightly older, more educated, less likely to be married, have fewer children, and are less likely to live in a multigenerational household. However, these are all aggregate trends that describe changes in Czech society over the past twenty years and seem to be unrelated to the fertility changes or parental allowance reforms.

3.6 Conclusion

This paper examines the impact of the duration of paid parental leave on the labor market status of women after childbirth, using two reforms that occurred in the Czech Republic in 1995 and 2008. We estimate the effect of the two reforms on the entire career break after childbirth, which we then disentangle into the effect on the family leave duration and the subsequent period of unemployment that many women experience when returning to the labor force.

We find that in response to the 1995 reform that extended the paid parental leave from three to four years, while keeping the job protection at three years, there is a 35 p.p. increase in inactivity of mothers whose youngest child is between 3 and 4. The rise in non–employment among these mothers is somewhat smaller due to the 10 p.p. fall in unemployment, as mothers who used to look for a job when their child was 3 years old prior to the reform now more often remain inactive for at least one more year. The 1995 reform increased both inactivity and unemployment of mothers with children aged 4 by about 6 and 5 p.p., respectively. Mothers of children aged 5 and 6 experienced an increase in unemployment probability by 4 p.p. and 2 p.p., but no impact on inactivity, and even this impact on unemployment faded away by the time the child turned 7. The reform therefore had no permanent impact on women's attachment to the labor force.[39]

[39]This is in contrast with in Schönberg and Ludsteck (2014), where a similar reform induces about 4%

A similar, but weaker, pattern in the opposite direction is obtained when estimating the impact of the 2008 reform, which allowed some mothers to shorten their leave from four years to three or two years, depending on their pre–birth income, while preserving the total amount of the allowance received. The flexible allowance schedule reduced the inactivity of women with 3 and 4 year old children by 21 p.p. and 7 p.p., respectively. It raised, however, the share of unemployed women with children of 2 and 3 by 3 p.p. and 7 p.p., as some mothers started looking for a job earlier than before the reform, due to the reduction in the duration of parental leave. As a result, the non–employment rate rose for mothers of 2 year olds and fell, but to a lesser extent, among mothers of children between 3 and 5, again with no effect on mothers of older children.

The absence of any effect of the reforms on mothers with children older than 7 (6) suggests that the impact of changes in paid parental leave duration on labor market status of women after childbirth disappears over time. Whether an impact persists in terms of the changes in post–birth earnings or job quality cannot be answered with our data and so is left for future research.

of women to remain out of the labor force by the time their child is 6.

3.A Appendix 3

Table 3.7: Difference–in–differences estimation for 1995 reform: Robustness check

	Treatment group: women with the youngest child					
	aged 2	*aged 3*	*aged 4*	*aged 5*	*aged 6*	*aged 7*
			Panel A: impact on non–employment			
Treat*After	0.032***	0.277***	0.135***	0.066***	0.038***	0.030***
	-0.007	-0.018	-0.016	-0.009	-0.006	-0.007
R-squared	0.337	0.174	0.084	0.068	0.064	0.064
Observations	145270	145424	145824	146012	146197	146125
			Panel B: impact on inactivity			
Treat*After	0.059***	0.367***	0.077***	0.018***	0.006	0.006
	-0.01	-0.018	-0.008	-0.007	-0.004	-0.005
R-squared	0.394	0.177	0.065	0.06	0.061	0.062
Observations	145270	145424	145824	146012	146197	146125
			Panel C: impact on unemployment			
Treat*After	-0.027***	-0.091***	0.058***	0.048***	0.032***	0.024***
	-0.004	-0.006	-0.013	-0.008	-0.006	-0.005
R-squared	0.016	0.026	0.042	0.03	0.024	0.022
Observations	145270	145424	145824	146012	146197	146125

Note: The treatment groups consist of prime–aged women (aged 25–55) whose youngest child is aged 2–7. The control group consists of prime–aged women whose youngest child is aged 13–23. For each treatment group, a separate regression is estimated by the age of the youngest child, but the control group is fixed in all regressions. All regressions include dummies for the treatment group and after period, quarter–year dummies, and other control variables. The after period is defined differently in each regression for both treatment and control groups according to the quarter in which children of women who were eligible for the post–reform parental allowance with this child reached that particular age. Standard errors (in parentheses) are clustered at the group–year level (* p<0.10, ** p<0.05, *** p<0.01). Source: Czech LFS (1993–1999), own calculations.

Table 3.8: Summary statistics by treatment group and period, 1995 reform

	Treatment group												Control group	
	Child aged 2		Child aged 3		Child aged 4		Child aged 5		Child aged 6		Child aged 7		Child aged 8-13	
	Before	After	Before	After	Before	After	Before	After	Before	After	Before	After	Before	After
Non–employed	0.854	0.906	0.377	0.677	0.211	0.359	0.162	0.243	0.144	0.186	0.126	0.175	0.091	0.13
Inactive	0.829	0.894	0.236	0.611	0.118	0.197	0.097	0.131	0.081	0.1	0.069	0.096	0.057	0.073
Unemployed	0.025	0.013	0.141	0.065	0.093	0.162	0.064	0.112	0.063	0.086	0.058	0.079	0.033	0.057
Age	29.98	29.98	30.54	30.39	31.10	31.16	31.64	31.82	33.00	32.57	33.85	33.62	37.67	37.37
Primary education	0.123	0.094	0.112	0.094	0.131	0.084	0.137	0.101	0.14	0.108	0.152	0.118	0.184	0.14
Secondary education	0.744	0.79	0.767	0.803	0.756	0.814	0.754	0.802	0.754	0.793	0.744	0.786	0.732	0.764
Tertiary education	0.133	0.116	0.121	0.103	0.113	0.101	0.109	0.097	0.107	0.099	0.104	0.096	0.084	0.096
Married	0.913	0.905	0.926	0.888	0.915	0.878	0.9	0.877	0.895	0.869	0.886	0.852	0.863	0.844
Cohabiting	0.034	0.042	0.022	0.041	0.025	0.043	0.033	0.039	0.032	0.035	0.031	0.04	0.033	0.037
Number of children	2.038	1.998	2.021	1.952	1.985	1.943	1.951	1.913	1.953	1.907	1.939	1.856	1.758	1.723
Presence of elderly	0.023	0.02	0.028	0.024	0.03	0.028	0.029	0.026	0.033	0.025	0.039	0.03	0.042	0.036
Observations	4654	6753	4578	6982	4756	7206	4795	7354	4852	7483	4622	7640	27802	44338

Notes: The treatment group consists of women whose youngest child is aged 2 to 7. Women whose youngest child is aged 8 to 13 serve as a control group. The reform was in force from October 1, 1995, so we define the before–reform period to cover two years and three quarters before the reform (1993 Q1–1995 Q3) and the after–reform period as three years and one quarter after the reform (1995 Q4–1998 Q4). Source: Czech LFS (1993–1998), own calculations.

Table 3.9: Summary statistics by treatment group and period, 2008 reform

	Treatment group												Control group	
	Child aged 2		Child aged 3		Child aged 4		Child aged 5		Child aged 6		Child aged 7		Child aged 8-13	
	Before	After	Before	After	Before	After	Before	After	Before	After	Before	After	Before	After
Non–employed	0.859	0.853	0.663	0.519	0.39	0.307	0.247	0.211	0.225	0.189	0.205	0.159	0.155	0.125
Inactive	0.85	0.84	0.619	0.43	0.208	0.148	0.095	0.09	0.091	0.079	0.077	0.07	0.066	0.057
Unemployed	0.009	0.013	0.044	0.089	0.182	0.159	0.152	0.121	0.134	0.111	0.128	0.089	0.088	0.067
Age	31.19	32.68	31.86	33.52	32.60	34.21	33.17	34.94	33.74	35.59	34.67	36.36	37.65	38.89
Primary education	0.067	0.051	0.066	0.068	0.075	0.071	0.076	0.077	0.089	0.081	0.084	0.066	0.09	0.072
Secondary education	0.786	0.728	0.816	0.75	0.803	0.767	0.793	0.782	0.807	0.774	0.808	0.804	0.81	0.795
Tertiary education	0.147	0.221	0.118	0.182	0.122	0.162	0.131	0.141	0.104	0.144	0.108	0.13	0.1	0.132
Married	0.829	0.773	0.839	0.76	0.816	0.764	0.803	0.732	0.792	0.741	0.766	0.75	0.764	0.738
Cohabiting	0.099	0.15	0.077	0.143	0.072	0.119	0.069	0.121	0.068	0.102	0.067	0.095	0.061	0.074
Number of children	1.83	1.769	1.845	1.821	1.859	1.863	1.809	1.826	1.828	1.794	1.79	1.76	1.66	1.596
Presence of elderly	0.016	0.02	0.014	0.018	0.017	0.023	0.025	0.028	0.028	0.026	0.026	0.026	0.029	0.031
Observations	6544	9335	6148	7690	5595	6741	5077	6285	5228	5884	5029	5542	32040	30849

Notes: The treatment group consists of women whose youngest child is aged 2 to 7. Women whose youngest child is aged 8-13 serve as a control group. The reform was in force from January 1, 2008, so we define the before–reform period to cover four years before the reform (2004 Q1–2007 Q4) and the after–reform period as four years after the reform (2008 Q1–2011 Q4). Source: Czech LFS (2004–2011), own calculations.

Table 3.10: Difference–in–differences estimation for 2008 reform: Robustness check

	Treatment group: women with the youngest child				
	aged 2	*aged 3*	*aged 4*	*aged 5*	*aged 6*
	Panel A: impact on non–employment				
Treat*After	0.023***	-0.154***	-0.090***	-0.01	-0.023**
	-0.005	-0.016	-0.01	-0.007	-0.01
R-squared	0.392	0.208	0.104	0.077	0.075
Observations	137741	135700	134198	133224	132974
	Panel B: impact on inactivity				
Treat*After	0.012**	-0.209***	-0.071***	0.009	-0.001
	-0.005	-0.02	-0.009	-0.006	-0.006
R-squared	0.505	0.237	0.053	0.036	0.036
Observations	137741	135700	134198	133224	132974
	Panel C: impact on unemployment				
Treat*After	0.011***	0.055***	-0.019**	-0.019***	-0.022***
	-0.003	-0.006	-0.008	-0.006	-0.006
R-squared	0.036	0.034	0.054	0.048	0.046
Observations	137741	135700	134198	133224	132974

Note: The treatment groups consist of prime–aged women (aged 25–55) whose youngest child is aged 2–6. The control group consists of prime–aged women whose youngest child is aged 18–25. For each treatment group, a separate regression is estimated by the age of the youngest child, but the control group is fixed in all regressions. All regressions include dummies for the treatment group and after period, quarter–year dummies, and other control variables. The after period is defined differently in each regression for both treatment and control groups according to the quarter in which children of women who were eligible for the post–reform parental allowance with this child reached that particular age. Standard errors (in parentheses) are clustered at the group–year level (* p<0.10, ** p<0.05, *** p<0.01). Source: Czech LFS (2004–2012), own calculations.

Table 3.11: Characteristics of Czech mothers with a new–born child

Year	Age	Primary education	Secondary education	Tertiary education	Married	Cohabiting	Number of children	Presence of elderly	Observations
1993	29.2	0.1	0.747	0.153	0.929	0.038	2.148	0.012	1200
1994	29.579	0.142	0.71	0.148	0.935	0.038	2.103	0.022	1078
1995	29.733	0.122	0.741	0.136	0.9	0.054	2.133	0.024	588
1996	29.467	0.085	0.753	0.162	0.912	0.058	2.167	0.023	1074
1997	29.466	0.099	0.727	0.174	0.887	0.065	2.133	0.02	999
1998	29.449	0.087	0.762	0.151	0.898	0.06	2.06	0.025	1063
1999	29.721	0.079	0.789	0.132	0.87	0.079	1.996	0.029	1050
2000	29.176	0.071	0.765	0.164	0.904	0.054	1.913	0.014	996
2001	29.454	0.075	0.76	0.165	0.853	0.078	1.903	0.021	1125
2002	29.382	0.076	0.762	0.162	0.858	0.074	1.865	0.014	1154
2003	30.157	0.061	0.786	0.154	0.823	0.102	1.875	0.014	1335
2004	30.183	0.076	0.765	0.16	0.83	0.11	1.88	0.02	1389
2005	30.254	0.068	0.746	0.186	0.819	0.133	1.829	0.017	1380
2006	30.402	0.052	0.751	0.197	0.808	0.143	1.818	0.009	1415
2007	30.809	0.055	0.752	0.194	0.771	0.157	1.837	0.008	1465
2008	31.363	0.063	0.723	0.214	0.731	0.196	1.827	0.01	1508
2009	31.387	0.058	0.72	0.222	0.739	0.191	1.844	0.013	1600
2010	31.365	0.057	0.676	0.268	0.701	0.227	1.819	0.012	1677
2011	31.641	0.052	0.613	0.335	0.73	0.206	1.791	0.017	1658
2012	31.674	0.046	0.619	0.335	0.715	0.211	1.768	0.016	1450

Notes: The table reports characteristics of women with children aged 0 from 1993 to 2012. The characteristics include average age, share of primary, secondary and tertiary educated, share of married and cohabiting, average number of children, share of women who live in a household with elderly (above 65) household members, and number of observations in the sample. Source: Czech LFS (1993–2012), own calculations.

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