Impact of Chilean Maternity Leave Expansion on Female Labor Market Outcomes & Gender Discrimination

Luis Faundez

Abstract

This paper uses a policy change in Chile, which in 2011 increased the paid maternity leave period from 18 weeks to 30 weeks, to study its impact on mothers of young children labor market attachment, and if this policy has any additional impact on labor market discrimination against women. Using a difference-in-differences approach, I find robust evidence that the aforementioned policy increased labor market attachment of mothers of infants. However, this policy seems to have had an unintended effect: it reduced labor force participation of women of childbearing age by 3 percentage points and their employment by 2.4 percentage points, while it had no effect on the gender pay gap.

JEL classification: J31, J71

I. Introduction

The effects of maternity leave benefits on female labor market outcomes such as participation, employment, and gender wage gap have been widely studied for developed countries, especially European countries. However, there is little evidence for less developed countries. These countries tend to have greater gender gaps in labor market outcomes such as employment rates and wages than developed countries. Part of these gender differences are attributed to the conventional social role of women as caregiver of children, which influences women's decision of participating in the labor force. In this paper, I study how increasing maternity leave benefits affects mothers of young children labor market attachment, and if this policy has any additional impact on labor market discrimination against women in Chile. In the case of Chile, it is particularly important to look at discrimination and inequality as it is the OECD country with the worst income distribution and the highest gender pay gap. This gap gets worse as women get more schooling: highly-educated women make 37 percentage points less than men, conditional on education level.

Maternity leave policies differ vastly in terms of duration of the benefit, job protection and income replacement across countries. The International Labour Organization standard is 14 weeks, so with the new policy, Chile is exceeding this standard by 16 weeks, but it is still below the OECD average of 52.6 weeks as of 2015. Among OECD countries, maternity leave benefits replace around 79% of previous earnings for a mother on average wages, with 13 OECD countries providing full earnings replacement, in most cases with a cap. Replacement rates tend to be lower in English-speaking countries, such as Australia, Canada, Ireland, New Zealand, and the United Kingdom, which replace less than 50% of previous earnings on average. Higher income mothers tend to have lower replacement rates due to the ceilings of the maternity leave benefits. For instance, in the Netherlands and Norway, the relatively low payment caps mean that replacement rates for a mother making 150% of average earnings are around 30- 40 percentage points lower than those for a mother on average earnings (OECD, 2016). Eastern European and

Nordic countries offer the most generous benefits: several months of paid leave, possibility of unpaid leave, and extended period of job protection.

The main objectives of maternity leave policies are to help mothers recover after giving birth and help them bond with their newborns, while providing mothers job security and income replacement. As Rossin-Slater, Ruhm, and Waldfogel (2013) point out, short maternity leaves are also intended to promote women's labor force attachment, while longer job absences may pose a risk to women's labor market position due to human capital depreciation and because employers may view women of childbearing age differently from other employees. Additionally, in Chile, policy maker's objectives include a reduction in uncertainty about when women would return to work, and sharing part of the costs of pregnancy from the firms' perspective between mothers' and fathers' employers, by entitling fathers with the possibility of sharing part of the maternity leave benefit.

This paper uses a policy change that in 2011 increased the paid maternity leave period from 18 weeks to 30 weeks, to study its impact on female labor force participation and employment rates, and on the gender pay gap. This study uses 6 waves from 2003 to 2015 of the CASEN survey from the Chilean Ministry of Social Development, and a difference-in-differences approach that compares pre- versus post-policy implementation labor market outcomes. In order to identify policy effects on mothers' labor market outcomes, I use women whose youngest child is younger than one year old as the treatment group. Additionally, I investigate if there is persistence of effects by using women whose youngest child is one year old, and women whose youngest child is two years old as treatment groups. These treatment groups' outcomes are compared to those outcomes of different control groups: women whose youngest child is between 5 and 10 years old, mothers whose youngest child is 11 to 18 years old, and childless women. The preferred control group is mother of younger children, as it is likely that the labor market for mother of infants (0 to 2 years old) is more similar to that of mothers of young kids (5 to 10 years old). Using the preferred approach, I find evidence that expanding maternity leave benefits effectively increased mothers of newborns' labor market attachment by increasing both employment and labor force participation by 5 percentage points, with not statistically significant impact on the gender pay gap. These effects decrease as their children grow older. These findings are in line with those of Rossin-Slater, Ruhm, and Waldfogel (2013), who found that mothers of 1- to 3-year old children increased their labor supply due to the implementation of California's paid family leave, and Gregg et al. (2007) who document that mother are returning to work within the first year after giving birth due to the introduction of maternity leave rights.

To assess the impact of the policy on labor market discrimination against women, the strategy is to compare outcomes of childbearing age women (18 to 38 years old), the treatment group, to older women (39 to 50 years old), the control group. In the preferred econometric specification, I find robust evidence that the 2011 maternity leave policy change reduced labor force participation of childbearing age women by 3 percentage points, while it had no effect on the gender pay gap. These findings contrast to those of Ruhm (1998), who found that paid parental leave mandates in nine European countries led to an increase in women's employment and a reduction in their relative wages.

This paper expands the literature on the impact of maternity leave policies on female labor market outcomes on several dimensions. First, it studies if a policy that intends to help and improve the labor market position of mothers of young children and women who desire to become mothers, has any impact on labor market attitudes towards women in general. Second, this study provides evidence for a less developed country of the impact of extending maternity leave benefits on women's labor force participation and employment rates, and the gender pay gap. Third, it provides evidence for a change in behavior induced only by the possibility of spending more time at home instead of working, as the job protection and income replacement aspects of the maternity leave coverage remained unchanged[1].

[1] Job protection starts from the moment the working woman gets pregnant and ends 15 months after giving birth. Income replacement was kept at 100% of previous mother's earnings with a cap, this ceiling is binding only for approximately 5% of working women in Chile as of June 2018, therefore it is safe to assume that there is no reduction in family income.

The remainder of this paper is structured as follows: section II discusses the conceptual framework, a literature review is provided in section III, section IV provides some background information on maternity leave policies in Chile, section V describes the data used in this study, section VI presents the empirical strategy to identify the policy effects, results and specification and robustness tests are shown in section VII, and section VIII concludes.

II. Conceptual framework

The effects of maternity leave rights on employment are theoretically ambiguous (Klerman and Leibowitz (1994)). After giving birth, a woman can be on a leave from her job or not employed. If the labor market is competitive, with perfect information and no externalities, maternity leave policies reduce economic efficiency by limiting the ability of employers and workers to voluntarily select the optimal compensation package (Ruhm (1998)).

Some argue that leave rights decrease female unemployment and increase firm-specific human capital by reducing the need for women to change jobs if they wish to spend more time at home with their newborns (Kamerman (1998)). With competitive labor markets, those more likely to use maternity leave will pay for it by receiving lower wages, which implies that women of childbearing age will get lower compensation if the benefit is mandated (Ruhm (1998)). Moreover, if the leave period is long enough, it may cause employers to limit women to positions where absences are less costly, increasing occupational segregation[2].

Under the presence of asymmetric information, an adverse selection problem might arise as a big proportion of “high-risk” childbearing-age women might self-select into companies that are required to or voluntarily provide more maternity benefits, which will force these companies to pay lower wages. On the other hand, “low-risk” women will sort out of these firms. As Ruhm (1998) points out, a government policy could potentially eliminate the incentives for this sorting and raise welfare.

[2] See Stoiber (1990) for some evidence on this issue in Sweden.

Maternity leave entitlements are likely to shift the labor supply curve to the right for those groups more likely to use the benefits. The demand curve would simultaneously shift to the left to the extent that non-wage costs increase[3], as the maternity leave benefits are paid by with public funds in Chile. It is important to mention that increase in costs would be partially offset by a reduction in employers' childcare expenses, as the Chilean legislation requires companies hiring twenty or more women to provide childcare for their female employees' children from the moment women return to work after maternity leave until the child turns 2 years old. Therefore, it can be expected that the shift in the labor supply is going to be larger than the shift in the labor demand, implying that the relative employment of childbearing age women will rise and their relative wages will fall in equilibrium. Employment might be reduced right after childbirth due to increased leave-taking; however, as Klerman and Leibowitz (1997) argue, employment may increase during this time period if some women who otherwise would have quitted their jobs to take more leave than previously allowed, now remain in their jobs and return to the workforce sooner than before.

There are some arguments for a shift to the right of the labor demand curve. There could be productivity gains if the maternity leave increases firm-specific human capital by allowing women to remain in their jobs. This demand shift will further increase employment and partially or completely offset the decline in wages (Ruhm (1998)).

III. Literature Review

Studies for Canada and Europe have consistently found that the take-up of paid maternity leaves is very high, close to universal (Baker and Milligan (2008), Dustmann and Schonberg (2012), Burgess et al. (2008)). In the case of Chile, the take-up is universal for all eligible women.

[3] Expenses associated with hiring and training temporary replacements, or efficiency losses in case of not hiring a replacement.

For the case of Europe, studies such as Ruhm (1998), Dustmann and Schonberg (2012), and Gregg et al. (2007) have found mixed effects of maternity leave policies. Ruhm (1998) studies the economic consequences of rights to paid parental leave in 9 European countries from 1969 to 1993. He finds that rights to paid leave raise the percentage of women employed (female employment-to-population ratio) between 3 and 4 percentage points, with a substantial effect for shorter durations of guaranteed work absence. Brief leave periods have small effect on women's earnings, but longer leaves are associated with a substantial reduction in relative wages of between 2 to 3 percentage points. Dustmann and Schonberg (2012) analyze different maternity leave policy changes in Germany, finding a considerable effect on mothers' labor supply in the short-run; however, they only find a small effect on the overall share of women who returned to the labor market in the long run, after the job protection period had expired. They also find no support that the expansions in leave coverage improved children's long-term outcomes, such as children's educational attainment, high track school attendance, and wages at the age of 28. Similarly, Dahl et al (2016) study the effects of a series of policy reforms that expanded paid maternity leave from 18 to 35 weeks in Norway. Using a regression discontinuity approach for each of the 6 reforms they evaluate, they find little effect on children's schooling, parental earnings and labor force participation, fertility and marriage. In the United Kingdom, Gregg et al. (2007) document that maternity rights have had a profound effect on employment, which varies by the wage opportunities of mothers. Maternity leave policies induce a behavioral change in when mothers return to work: several of who previously would have not have gone back to work until their children were 3-5 years old are now returning to work within the first year. Their evidence suggests that this effect is most marked among better educated and higher earning mothers.

Labor markets tend to punish maternity. As Waldfogel (1998) points out, there is a "family gap" between the wages of mothers and non-mothers in the United States and the UK. About 40% to 50% of the gender gap is explained by differential returns to marital and parental status. Women's lower level of experience and lower returns to experience explain another 30% to 40% of the gap. Maternity leave policies affect mainly younger women in fertile age; hence, such policies might have a role to play in explaining gender pay gap. In her research, Waldfogel (1998) finds that women who had maternity leave coverage and returned to work after childbirth received a wage premium that offset the negative wage effects of children, suggesting that maternity leave reduces the family gap. Similarly, Baker and Milligan (2008), using Canadian data, find that the introduction of modest leaves of 17-18 weeks increases the proportion of mothers employed and on leave, but has little effect on the length of time they spend at home with their children.

Turning the attention to the United States, Rossin-Slater, Ruhm, and Waldfogel (2013) studied the effects of California's Paid Family Leave program, which took effect in 2004, on mothers' leave taking and labor market outcomes. They found evidence that the California program doubled the overall use of maternity leave, with some evidence of particularly large growth for less advantage groups, such as unmarried women, black and Hispanic women. Their findings also suggest that the program increased the usual weekly work hours of employed mothers of 1 to 3-year-old children by 10% to 17%.

Low and Sanchez-Marcos (2015) use a life-cycle model of female labor supply and savings behavior calibrated to the US economy to study the effect of introducing a maternity policy similar to Scandinavian-type policies on gender differences in participation rates and wages. They distinguish between the effect of the job protection offered by maternity leave and the effect of income replacement, finding that job protection leads to increase in participation of mothers with children under 6 years old, and minimal effects on wages, with the negative selection effects offsetting the reduced human capital depreciation. On the other hand, they found that the income replacement effect was limited on participation and wages

IV. Institutional background

Chilean working women have had maternity benefits for over 100 years now[4]. It was in 1917 when the first benefit was introduced. By then, those companies employing 50 o more women 18 or older needed to provide a space in their premises where women could leave their children younger than 1 year old, and at this place, women could also breastfeed their children for up to one hour a day. This time feeding their babies had to be paid. Later, in 1925 the first maternity leave benefit was enacted. This benefit consisted of 60 paid days of leave, 40 days before giving birth and 20 days after giving birth. Women were entitled to receive 50% of their salary, and it was completely paid by their employers. Years later, in 1931, the leave period was extended to a total of 12 weeks and the cost of the leave was now shared by the employers and the social security system. By 1952, income replacement was increased to 100% and was fully paid by the social security system.

Right before the policy change studied in this paper, the benefits included 6 weeks of paid prenatal leave, 12 weeks of paid postnatal leave, and paid sick leave in case of serious illness of children younger than one year old. In October of 2011 there was a modification to the law governing maternity leaves, which extended the benefit for working mothers. The new law allows women to have extra 12 weeks of paid postnatal maternity leave if they choose to absent from work full-time, or 18 extra weeks if they choose to be away from work part-time (half of their regular working time). The legislation also introduced a third option, women can transfer part of the maternity leave to the working father of the child after the sixth week of the new leave period for the number of weeks that the mother indicates, for up to 6 weeks of full-time leave or 12 weeks of part-time leave. In this last case, the amount of the subsidy is determined according to the father's salary. This benefit also applies to self-employed mothers that are part of the social security system.

[4] Romanik (2014).

The subsidy is for the full salary amount of the beneficiary and it is paid by with public funds from Fondo Unico de Prestaciones Familiares y Subsidios de Cesantia, which is financed with general taxes. The amount of the subsidy is capped at the maximum amount used to determine the contribution to the social security system[5]. This provides lower benefits for highest earning women. Eligibility requirements are easily met: women need to have had contributed to the social security system for the first time at least 6 months before the beginning of the prenatal leave and had been working for three continuous months before the beginning of the prenatal leave.

The Chilean legislation includes additional benefits for women during pregnancy and after giving birth. A pregnant woman cannot be fired during her pregnancy and her job is also protected for one year after the end of the first 12 weeks of the postnatal leave. Once the mother has come back to work, she has the right to absent from work for one hour a day in order to feed her newborn until the child turns 2. Also, those firms that employ 20 or more women are required to provide childcare to their workers' children until they turn 2. These benefits impose higher costs of hiring women, especially in fertile age, which should explain part of the gender pay gap. According to OECD statistics, Chile is the OECD country with the highest mean gender pay gap conditional on education level, the pay gap among full-time-employed highly-educated men and women is over 37 percentage points, while the gap for full-time-employed low-skilled workers is around 23 percentage points.

One of the objectives of the extension of the maternity leave benefit was to reduce the costs associated with hiring women. This new benefit was expected to reduce uncertainty regarding when women will reincorporate to work, as it was a common practice to use the paid sick leave in case of serious illness of children younger than one year old. Also, there is a reduction in childcare costs for employers. Now employers have to finance 12 weeks less of childcare. Part of the cost of having children could be assume by fathers' employers, if men make use of the leave.

[5] Approximately USD $3,300 per month.

Hence, hiring women should be relatively less expensive compared to men than before. However, men are not using the benefit, as it can be seen on figure 1.

Since its implementation in October 17, 2011, the majority of women have chosen the full-time option of the new benefit. When the new policy was just implemented, 7% of mothers chose the part-time option, then this proportion started to decrease and stabilized at around 1% in 2013. The relatively high proportion of the part-time option usage at the beginning of the new policy is explained by women who had already given birth when the policy took effect, but still met the conditions set for using the new benefit. Since these women had already agreed with their employers to come back to work after 12 weeks, many of them opted for using the part-time option of the new benefit. The proportion of men using the benefit has remained relatively constant at around 0.3%. Figure 1 shows men's usage of the extra 12 weeks of maternity leave benefit since its implementation in 2011.

V. The data

For this paper, I use repeated cross-sectional survey data from the Chilean Ministry of Social Development. The CASEN survey (Encuesta de Caracterización Socioeconómica Nacional) is conducted usually every other year, and consists of a representative sample of households of the whole country. It collects individual data on demographic characteristics, education, employment, income, health, and housing variables. I use the 2003, 2006, 2009, 2011, 2013, and 2015 waves.

My sample consists of 721,055 women and men between 18 and 50 years old. Women account for 51.51% of the sample. Table 1 shows a summary of descriptive statistics of the sample by gender and cohort in 2009. The younger cohort consist of people between 18 and 38 years old, and the older cohort comprises people between 39 and 50 years old. Women are a bit older than men, on average, more likely to be single (except younger women), and more likely to live in an urban area. Women also have a higher level of schooling, 10.2% of younger women have a college degree versus only 7.9% of younger men. Among the older cohort, gender college attainment difference is only 1 percentage point. In the same way, the greatest difference in schooling is across cohorts, younger men and women get 2 more years of schooling on average than their older counterparts. Despite this big difference in schooling across cohorts, older women make 9.5% more than the younger cohort and this cohort difference for men is 28.1%, suggesting that experience and/or tenure play an important role determining earnings. Surprisingly, despite the fact that women are more educated than men, they make significantly less than men (25.26% less on average). The unconditional gender pay gap is larger for the older cohort than the younger cohort, 31 percentage points versus 19 percentage points, respectively.

Analyzing the labor force participation rate by gender, I find that men's participation rate slightly declined until 2009 and from there onwards has remained stable around 83%. Female participation rate, on the other hand, has been continuously increasing, with a hike of almost 10 percentage points between 2009 and 2011. By 2015, women's participation was close to 60%. Figure 3 depicts the participation rates by gender for the sample period. Similarly, employment-to-population ratios by gender follow the same pattern, which is shown in Figure 5.

There is only a 1.5 percentage points difference in labor force participation rate between childbearing age women and their older counterpart, while young women's participation is 48.3%, older women's participation is almost 50%. However, difference in employment-to-population ratio between treatment and control group is bigger: 5.8 percentage points higher for older women.

Younger women have slightly larger households than older women. While women between 39 and 50 years old have 1.64 children living in the household, childbearing women have only 0.74 children on average. Analyzing the age of the children in the household, unsurprisingly 19% of younger women have at least one child who is 4 or younger versus 7% of older women. Similarly, 11.5% of childbearing age women have at least one child who is 2 or younger, while only 3% of older women do. Finally, looking at infants in the household, Figure 2 shows that the proportion of women who has a child younger than one year old starts to drop after turning 30 years old, and by age 39 only 2% of women are mother of an infant, by age 45 almost no woman has recently giving birth. This provides some support to the validity of choosing the older women group as the control group for the labor market discrimination question.

VI. Empirical Strategy

Below I describe the identification strategies to answer both research questions: the effects of expanding maternity leave benefits on labor market outcomes of the affected group, mothers of infants, and the effects on women of childbearing age in general, as a proxy for discrimination.

A. Labor market discrimination

I use a differences-in-differences approach that compares changes in labor market outcomes for the treatment group to changes in outcomes for the control group before and after the policy change in 2011. Since maternity leave policies are more likely to affect only women in childbearing age, the treatment group is composed of women between 18 and 38 years old. As a control group, I use older women between 39 and 50 years old, since they are less likely to get pregnant and hence should be viewed by employers as unlikely to use the policy benefits[6].

For employment rate and labor force participation, I estimate the following equation:

$$y_{it} = \beta_0 + \beta_1 Treatment_i \times Post_t + \gamma_t + \zeta_a + \mathrm{X}'\delta + \varepsilon_{it} \qquad (1)$$

Where y_{it} is a dummy variable for the relevant outcome for individual *i* surveyed in year t, Treatment is a dummy indicating being 18 to 38 years old, Post is a dummy equal to 1 for 2011 onwards, γ_t represents year fixed effects, ζ_a are age dummies, and X is a vector of covariates such as schooling, marital status, number of children, etc. The coefficient of interest is β_1, which

[6] Even though it is biologically possible for a woman to get naturally pregnant after age 40 and there are higher chances with assisted methods (IVF), there seems to be a social age deadline for childbearing of women, which is perhaps more important. Billari et al (2011) show that across 25 European countries a maternal social age deadline of ≤40 years of age is perceived for the majority of the population.

will tell us the change in employment rate or participation rate for childbearing age women that is attributable to the maternity leave policy.

As a way to test for differential pre-trends and examine how the treatment effect varies across time, I estimate the following variation of equation 1:

$$y_{it} = \sum_{k=1}^{6} \beta_k Treatment_i \times Year_t + \gamma_t + \zeta_a + X'\delta + \varepsilon_{it} \qquad (1a)$$

Each β_k coefficient represent the treatment effect for each of the 6 years of data used in this study. If the common trend assumption is plausible, then all β_k until 2009 should be zero.

To study the effects of the maternity leave extension on the gender pay gap, I estimate the following equation:

$$\begin{aligned} y_{it} = \beta_0 + \beta_1 Female_i + \beta_2 Treat_i \times Post_t \times Female_i + \beta_3 Treat_i \times Female_i \\ + \beta_4 Post_t \times Female_i + \gamma_t + \zeta_a + \pi_o + X'\delta + \varepsilon_{it} \qquad (2) \end{aligned}$$

Where y_{it} is the log of monthly labor income for individual *i* surveyed in year *t*, Treat is a dummy indicating being 18 to 38 years old, Post is a dummy equal to 1 for 2011 onwards, Female is dummy variable for gender, γ_t represent year fixed effects, ζ_a are age dummies, π_o represents occupation fixed effects, and X is a vector of covariates such as schooling, experience, etc.

For this outcome, I also estimate a variation of equation 2 (equation 2a) in order to test for pre-trends and differential treatment effects over time. Each β_k coefficient before the year of implementation of the policy, 2011, is expected to be zero.

$$\begin{aligned} y_{it} = \sum_{k=1}^{6} \beta_k Treat_i \times Year_t \times Female_i + \beta_7 Female_i + \beta_8 Treat_i \times Female_i \\ + \sum_{j=1}^{6} \beta_j Year_t \times Female_i + \gamma_t + \zeta_a + \pi_o + X'\delta + \varepsilon_{it} \qquad (2a) \end{aligned}$$

One condition to get consistent estimates with a differences-in-differences strategy is the common trend assumption. In this case, the identification assumption is that absent this maternity leave policy change, outcomes for childbearing age women and older women would have trended in the same way. Figure 4 shows labor force participation rates since 2003 until 2015 for both groups. It can be seen that before the policy change, both groups followed a similar upward trend on the participation rate, making plausible the common trend assumption. Between 2009 and 2011 there was a big jump in the participation rate for both groups[7]. Employment-to-population ratio followed a similar pattern, which is shown in Figure 6. Figure 7 show the estimated conditional gender pay gap for childbearing age women and for older women. Gender pay gap for both groups was trending in the same way before the policy change, which gives support to the plausibility of the common trend assumption for this outcome.

B. Effect on mothers' labor market outcomes

To estimate the effects of the increase in maternity leave period on women actually targeted by the policy, I use a similar differences-in-differences approach. Such approach compares labor market outcomes of women having a child younger than one year old, the treatment group, to outcomes of mothers of older children, the control group. As before, I use equation 1 to estimate impacts on employment and labor force participation, and equation 2 to estimate the effect on the gender pay gap. Variables in both equations are defined in the same way as in the previous research question, except for the treatment variable, which now takes value 1 if individual *i* is a mother of a child younger than one year old, and zero if such individual is a mother whose youngest child is between 5 and 10 years old.

[7] This increase coincides with the economic recovery after the 2009 recession and the reconstruction work after the big earthquake of February 2010. During this period there was a generalized increase in employment for both men and women.

VII. Results

A. Labor market discrimination

In this section I investigate how the extension of 12 weeks in paid postnatal leave affected employment, labor force participation and the conditional gender pay gap for childbearing age women. I first look at female employment and labor force participation effects. One key concern with difference-in-difference designs is the common trend assumption; hence, my first approach is to do a detailed event study which will tell me about the pre-trends and the behavior of the treatment effect over time. Figure 8 shows the estimates for the policy effect on employment-to-population ratios by year, and Figure 9 shows the estimates for labor force participation rates by year. As expected, “treatment” effects before the policy took place in 2011 are statistically indistinguishable from zero, which suggests that both groups’ unconditional employment and labor force participation rates were trending parallel. This provides strong support for the validity of the research design. From 2011 onwards, the effect of the policy on both outcomes is negative and it has been increasing over time. Then, I estimate equation 1, which results are shown in columns 1 and 2 of Table 2. The maternity leave extension policy decreased childbearing age women’s unconditional employment rate by 2.4 percentage points and decreased female labor force participation rate by 3 percentage points. These results are significant at the 0.1% level.

Given that women already had maternity leave benefits, it is likely that women who want to work and become mothers in the near future were already participating in the labor force; thus, this postnatal leave extension only induced a small number of women to enter the labor force, if any. Although one could expect a very small response of the labor demand to this new policy, as non-wage costs of having newly-mother employees absent from work for 12 extra weeks should be small[8], the result suggests that labor demand shifted more than labor supply did.

[8] If a company was going to replace a mother-to-be employee before for only 18 weeks, it is likely that with the new policy the company was still going to replace the employee and just keep the temporary worker for the 12 extra weeks. Now, if a company would have not replaced the employee before, probably with the new policy is more likely to employ a replacement as the total absence would be for over 7 months.

Second, I proceed to study the effects on the gender pay gap. As with the previous outcome, my first approach is to conduct an event study. Figure 10 shows the estimates of the maternity leave extension on the conditional gender pay gap. Except for 2003, the estimated treatment effects are all indistinguishable from zero. In this case the evidence for the validity of the research design is not as strong as in the labor force participation rate case, but it is still plausible that the gender pay gap for both groups was trending in a similar fashion before the policy change. Moving to the regression framework, I estimate equation 2. Results are shown in column 3 of Table 2. This maternity leave extension had no impact on the gender pay gap for childbearing age women.

This result is somewhat surprising. A policy like this that intended to balance costs of hiring women and men by entitling men with parental leave rights should reduce the gender pay gap for childbearing age women. However, as Figure 1 shows, men are not using this benefit; hence, one could expect that hiring childbearing age women would become more expensive relative to hiring older women compared to men, increasing the gender pay gap. This null effect on the gender pay gap is not consistent with statistical discrimination against women. Employers discriminate in the same way younger women and older women compared to men, even though hiring younger women is relatively more expensive.

B. Effects on mothers' labor market outcomes

In order to answer the question about the effects of the maternity leave expansion on mothers, I estimate equations 1 and 2 using mothers of infants as the treatment group. My first approach uses women having a child younger than 1 year old as the treatment group and women whose youngest child is between 5 and 10 years old as the control group. The differences-in-differences estimates suggest that this policy increased mothers of infants' labor force participation by 5.8 percentage points and employment by 5.5 percentage points, while it had no statistically significant impact on the gender pay gap. These results are shown in Table 3. I conduct an event study for each outcome to test the plausibility of the common trend assumption. Figure 11 plots the estimates for unconditional employment, which suggest that the common trend assumption might hold. The graph also shows that the increase in employment peaked in 2013. Regarding labor force participation, Figure 12 depicts a similar trend as employment, and in this case the policy effect estimates for the pre-intervention period are zero. The event study for the conditional gender pay gap, presented in Figure 13, shows that only the estimate for 2003 is statistically different from zero.

Then I investigate if these effects on employment and participation lingers over time. To do so, I estimate equation 1 using different treatment groups. First, I use mothers whose youngest child is 1 year old as the treatment group and mothers whose youngest child is between 5 and 10 years old. Since a one-year-old child in 2011 was born in 2010, observations from this year are considered as part of the pre-intervention period. The differences-in-differences estimates shows that this increase in post-natal maternity leave increased mothers' employment and labor force participation by 2.4 and 2.9 percentage points, respectively, when their children were one year old. Again, I find no effect on the gender pay gap by estimating equation 2. My second approach is to use mothers whose youngest child is 2 years old as the treatment group and the same control group as before. This approach shows an increase of 2 percentage points in labor force participation and 1 percentage point increase in employment, however this estimate is not statistically significant. This approach also finds no statistically significant effect on the gender pay gap.

I also examine how the impact on employment and labor force participation varies according to marital status. I would expect a small or even zero effect for single mothers, as it is likely that most single women would have worked in the absence of this policy. Surprisingly, both employment and participation increases are larger for single mothers: 8.8 and 10.7 percentage points, respectively, versus 5.3 percentage increase in both outcomes for mothers living with their partner. These estimates are presented in table 10.

C. Specification and robustness tests

i) Labor market discrimination

A differences-in-differences design does not require any additional control if the identifying assumptions hold, in this case that labor market outcomes for childbearing age women and older women would have trended in the same way absent this maternity leave policy. I start by running these simple regressions for each outcome, which are shown in column 1 of tables 11, 12 and 13, for employment, labor force participation, and gender pay gap, respectively. These regressions would give an unbiased estimate of the policy effect if there are not unobserved characteristics that would have made outcomes of both groups trend differentially. There are some reasons for which this assumption might not hold. As people age, they might have different preferences about working, which might be also influenced by the general conditions of the labor market at different points in time. To control for unobserved differences in labor market attachment across the life path of people, I include age dummies. To eliminate any possible bias arising from unobserved conditions affecting the labor market in a particular year, I include year fixed effects. Column 2 of each specification table reports the estimates that include both year fixed effects and age dummies. Another threat to identification is that factors that might determine employment and wages, such as education attainment might vary differentially across younger and older women, as general educational attainment has been rising over time in Chile. To address this last point, I include covariates such as schooling, marital status and whether the individual is the head of the household to the regression.

For the case of employment-to-population ratio, the estimated effect of the policy using the simple DID regression is a reduction of 3.96 percentage points in childbearing age women's employment rate. Column 2 of table 11 suggest that there are differential trends across younger and older women due to age or time, as the point estimate remains unchanged. However, covariates seem to play a role, when they are added instead of fixed effects, I find a smaller effect of the policy, only 2.46 percentage points reduction in employment rate. Column 4 shows the estimates of a regression that adds geographical dummies to the main specification. The point estimate is similar to column 3, 2.69 percentage points reduction in female employment.

For the case of female labor force participation, the simple DID model (column 1 of table 12) gives an estimated effect of the policy of 4.4 percentage points reduction in the participation rate of childbearing age women. When year fixed effects and age dummies are added (column 2), I find a similar decrease of 4.65 percentage points in labor force participation. Column 3 shows the estimates of a regression that instead of fixed effects adds controls. Now I find a smaller effect of the policy, only 2.94 percentage points reduction in participation rate. Finally, column 4 adds geographic dummies to the main specification, finding a larger reduction in childbearing age women labor force participation of 3.4 percentage points.

I repeat the above exercise for the gender pay gap estimates (equation 2). Results for these regressions are shown in table 13. Column 1 includes no year fixed effect or age dummies, and controls flexibly for schooling (schooling dummies). Column 2 controls linearly for schooling instead and includes year fixed effects and age dummies. The point estimates in both columns are small, negative and not statistically different from zero. Column 3 also controls linearly for schooling but includes no year fixed effects or age dummies. Column 4 includes both year fixed effects plus age dummies, controls flexibly for schooling, and adds geographical fixed effects. Point estimates for these last two columns of table 13 are similar to the main results presented earlier, always small and statistically insignificant.

Overall, results for all outcomes are robust to different specifications. Thus, the evidence for the maternity leave extension of 12 weeks in Chile studied here suggests that the policy reduced employment and labor force participation for childbearing age women by 2.4 and 3 percentage points, respectively, and had no impact on the gender pay gap for younger women.

i) Effects on mothers of infants' labor market outcomes

I estimate equations 1 and 2 using alternative control groups: women whose youngest child is between 11 to 18 years old and childless women. Using the first control group, I find large effects on both employment and labor force participation of 5.6 and 6.2 percentage points, respectively. Using childless women as the control group, I find even larger effects on both outcomes: increase of 8.9 percentage points in employment and an increase of 9.9 percentage points in labor force participation. Both control groups give statistically insignificant increases in the gender pay gap between 3 and 4.7 percentage points. Tables 4 and 5 display these estimates. The main estimates are somewhat sensitive in magnitude to different control groups; however, when the control group is also composed by women with children (older kids), the point estimates are close to the main specification. This suggest that the best comparison group is women with children, as it is more likely that they and mothers of infants face a similar labor market, while the market for childless women is likely to be different from that of women with a baby.

Additionally, I conduct a falsification test. I estimate equations 1 and 2 using women whose youngest child is between 5 and 10 years old, the control group in the main approach of this research question, as the treatment group, and compare their outcomes to two different control groups. The first control group consists of women whose youngest child is between 11 and 15 years old. Here, I find no effect on any of the outcomes studied in this paper. Then, I expand this control group to include all women whose youngest child is between 11 and 18 years old. Similarly, I find no effect on any of the outcomes. Estimates for these two approaches are reported in tables 8 and 9. As expected, I find no effect on this group of women who should not have been affected by the expansion of the maternity leave period in 2011 as their youngest child was born in 2010, at the latest.

These results suggest that the 2011 maternity leave expansion in Chile increased labor market attachment of the targeted population: women with babies. Labor force participation and employment increased by at least 5 percentage points. Estimates of changes in gender pay gap are not precise, but they suggest a possible increase of it of around 3 percentage points.

VIII. Conclusions

This paper studies the impact of a maternity leave policy that in 2011 increased the paid postnatal leave period by 12 weeks in Chile. I find robust evidence that this policy increased labor market attachment of women with infants. This effect is greater the younger the child: mothers of a child younger than one year old increased their employment by 5.5 percentage points, when the child is one year old, the increase in employment is 2.3 percentage points, and when the kid is 2 years old, the increase is only 1 percentage point, although it is not statistically significant. Gender pay gap estimates are not statistically different from zero, but they do suggest a worsening of mothers' relative labor income. However, this policy seems to have had an unintended effect: it reduced labor force participation of women of childbearing age by 3 percentage points and their employment by 2.4 percentage points, while it had no effect on the gender pay gap. These last results suggest that probably the shift to the right of the labor supply curve was smaller than the shift to the left of the labor demand, employers are more responsive to this policy change than childbearing age women are. These findings are somewhat opposite to previous work that found that paid parental leave mandates in European countries led to increase in women's employment and a reduction in their relative wages. A policy that intended to improve mothers' position in the labor market, by allowing them to remain attached to the labor market, and help balancing the costs of hiring women and men by entitling men with parental leave rights, as well as improve newborns and their mothers' quality of life, had a negative impact on childbearing age women's employment and did not help reducing the gender wage inequality. This last point suggests that a public policy that could potentially reduce the gender pay gap is one that gives more mandatory benefits to fathers, as in the current policy is the mother's decision to transfer part of the leave period to the newborn's working father. Overall, these results are consistent with a scenario where women already in the labor force are incentivized to remain attached to it after giving birth, but other childbearing age women willing to enter the labor force are having difficulties finding a job, as hiring them is relatively more expensive for employers than hiring men or older women.

References

Baker, Michael, and Kevin Milligan. 2008. “How Does Job-Protected Maternity Leave Affect Mothers’ Employment?”. Journal of Labor Economics, Vol. 26, No. 4 (October 2008), pp. 655-691

Bill number 20545 (Ley 20545). 2011. “Modifica las normas sobre protección a la maternidad e incorpora el permiso postnatal parental”. http://www.leychile.cl

Billari, F.C., A. Goisis, A.C. Liefbroer, R.A. Settersten, A. Aassve, G. Hagestad, and Z. Spéder. 2011. “Social age deadlines for the childbearing of women and men”. Human Reproduction, Vol. 26, Issue 3, 1 March 2011, pp 616–622.

Burgess, Simon, Paul Gregg, Carol Propper, and Elizabeth Washbrook. 2008. Labour Economics, 15, pp. 168-201.

Dahl, Gordon, Katrine V. Løken, Magne Mogstad, and Kari Vea Salvanes. 2016. “What Is the Case for Paid Maternity Leave?”. The Review of Economics and Statistics, 2016, 98:4, pp. 655-670.

Dustmann, Christian, and Uta Schonberg. 2012. “Expansions in maternity leave coverage and Children’s long-term outcomes”. American Economic Journal: Applied Economics, 2012, Vol. 4, No. 3, pp. 190-224.

Gregg, Paul, Maria Gutierrez-Domenech, and Jane Waldfogel. 2007. “The employment of married mothers in Great Britain, 1974-2000”. Economica, 74, pp. 842-864.

International Labour Organization’s standards on Maternity protection. http://www.ilo.org

Kamerman, Sheila B. 1988. "Maternity and parenting benefits: An international overview." Zigler and Frank The Parental Leave Crisis, pp. 235-244.

Klerman, Jacob, and Arleen Leibowitz. 1994. “The Work-Employment Distinction among New Mothers”. The Journal of Human Resources, Vol. 29, No. 2, pp. 277-303.

Low, Hamish, Viginia Sanchez-Marcos. 2015. “Female labour market outcomes and the impact of maternity leave policies”. IZA Journal of Labor Economics.

OECD. Gender pay gaps for full-time workers and earnings differentials by educational attainment.
http://www.oecd.org/els/family/LMF_1_5_Gender_pay_gaps_for_full_time_workers.pdf

OECD. Parental leave replacement rates. December 2016. http://www.oecd.org/els/family/PF2_4_Parental_leave_replacement_rates.pdf

OECD Statistics on Length of maternity leave, parental leave and paid father-specific leave. http://www.oecd.org

Romanik, Katy. 2014. "Evolución de la legislación sobre protección a la maternidad". Cuaderno de Investigación N 58, Dirección del Trabajo, Gobierno de Chile.

Rossin-Slater, Maya, Christopher Ruhm, and Jane Waldfogel. 2013. "The effects of California's Paid Family Leave program on mothers' leave-taking and subsequent labor market outcomes". Journal of Policy Analysis and Management, Vol. 32, No. 2, pp. 224-245.

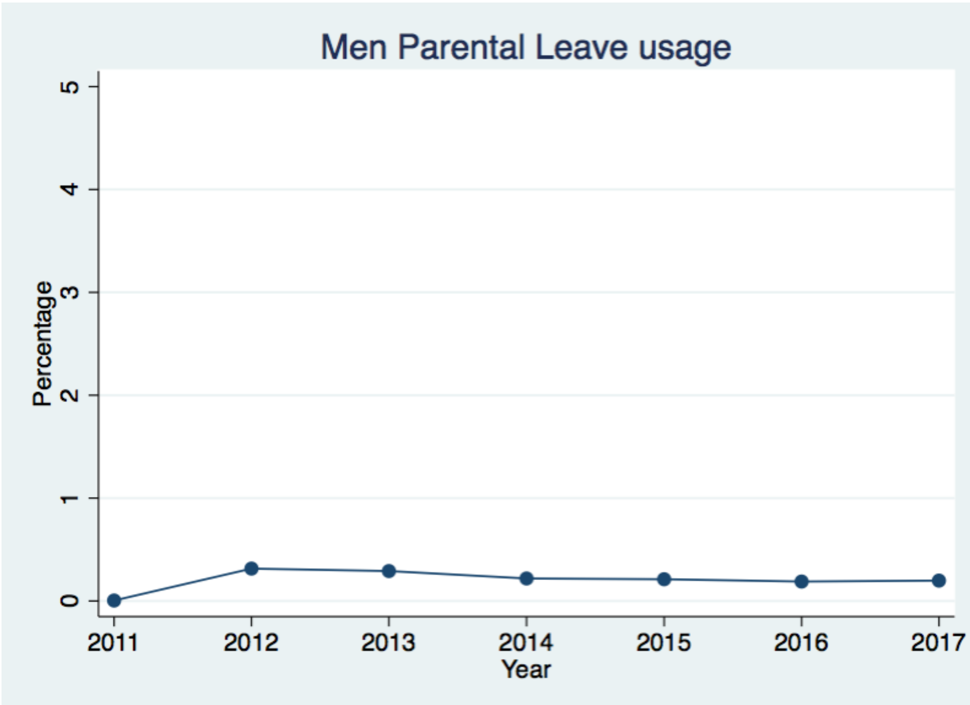
Ruhm, Christopher. 1998. "The economic consequences of parental leave mandates: Lessons from Europe". Quarterly Journal of Economics, 113, pp. 285-317.

Stoiber, Susanne. 1990. "Family Leave Entitlements in Europe: Lessons for the United States". Compensation & Benefits Management, Vol. 6, No. 2.

Waldfogel, Jane. 1998. "The Family Gap for Young Women in the United States and Britain: Can Maternity Leave Make a Difference?". Journal of Labor Economics, Vol. 16, No. 3 (July 1998), pp. 505-545

Tables and Figures

Figure 1: Parental Leave usage by men



Note: Proportion of eligible men taking the parental leave.

Table 1: Descriptive statistics by gender and age group

	Older Men	***Younger Men***	***Older women***	***Younger women***
Age	44.6	26.9	44.5	27.2
Schooling	9.16	11.15	9.31	11.41
College degree	7.6	7.9	8.5	10.2
Single	25.1	63.5	28.3	55.3
Urban	61.8	65.5	65.7	66.9
People in household	4.2	4.5	4.3	4.6
Head of household	71	26.8	22.6	9.6
Number of children	1.47	0.46	1.64	0.74
Have child 4 and younger	10.9	13.9	7	18.7
Have child 2 and younger	5.8	8.7	3	11.5
Monthly labor income[9]	506,812	395,630	349,772	319,272
Labor force participation rate	92.1	77.9	49.8	48.3
Employment rate	87.2	68.2	45.5	39.7
Observations	120,665	228,933	132,998	238,459

[9] Figures in Chilean pesos (CLP). Approximately 600 CLP = 1 USD.

Figure 2: Proportion of women who has a child younger than 1 year old by age

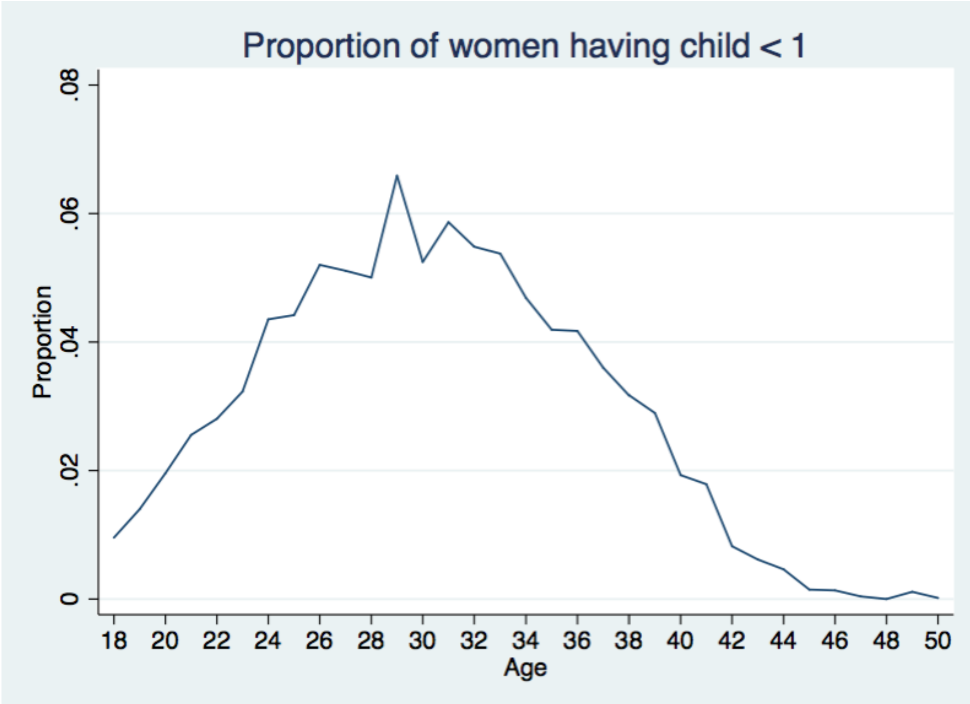


Figure 3: Labor force participation rates by gender and year

Figure 4: Female labor force participation Rates by age group and year

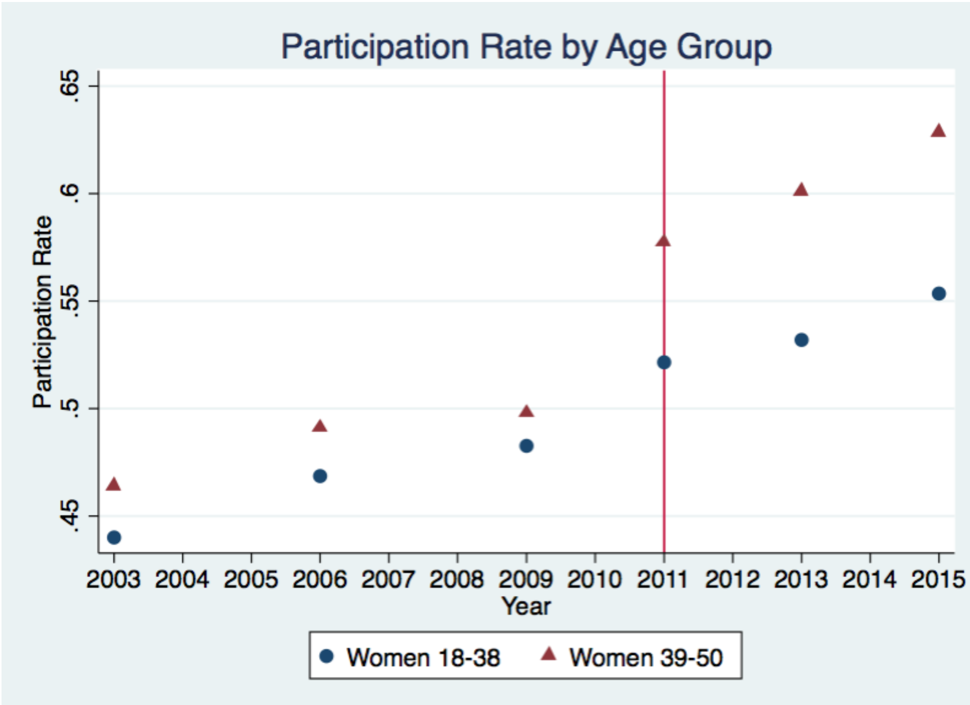


Figure 5: Employment-to-population ratios by gender and year

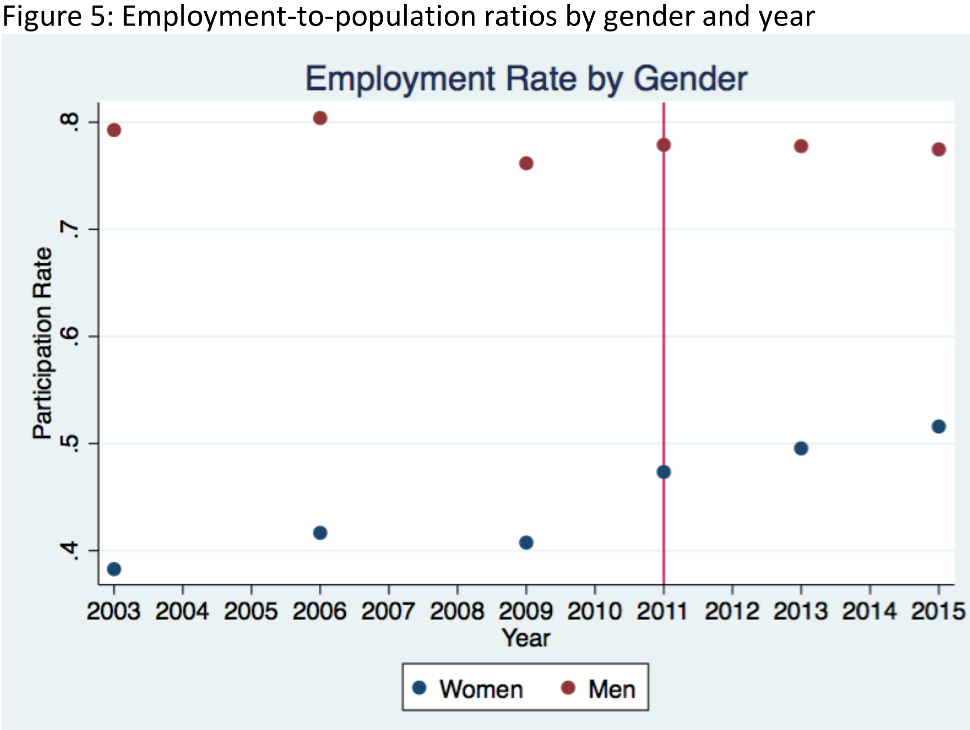


Figure 6: Female employment-to-population ratios by age group and year

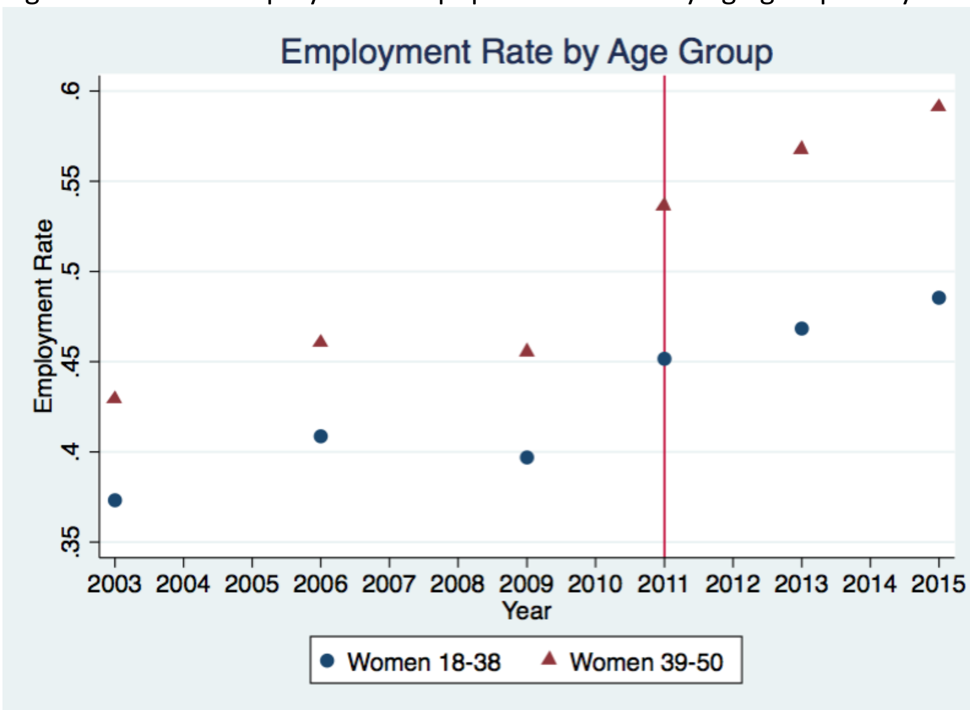
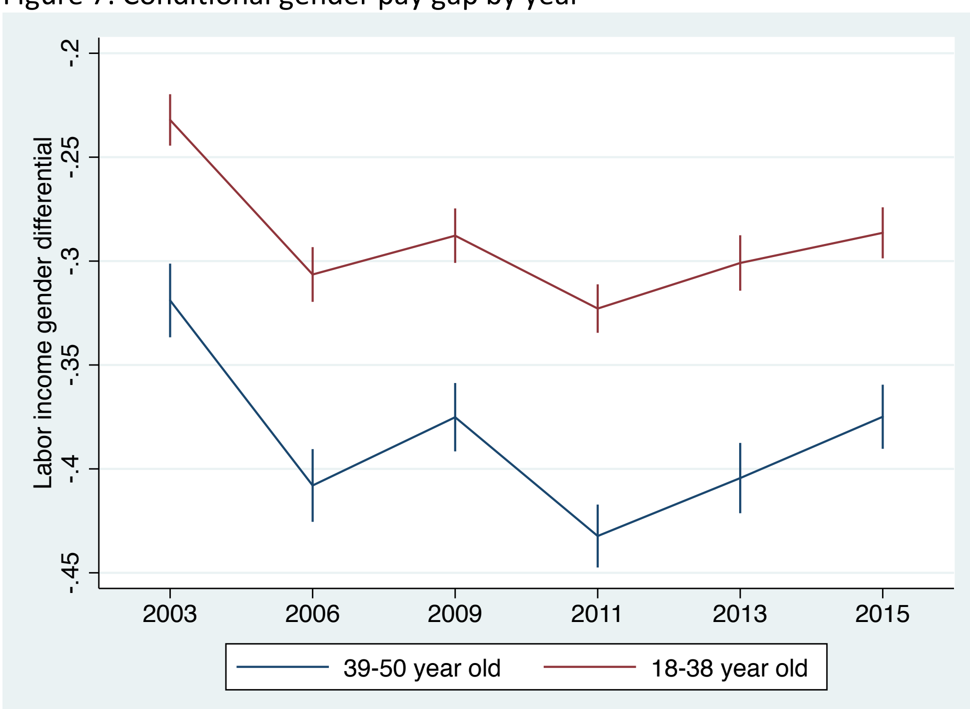


Figure 7: Conditional gender pay gap by year



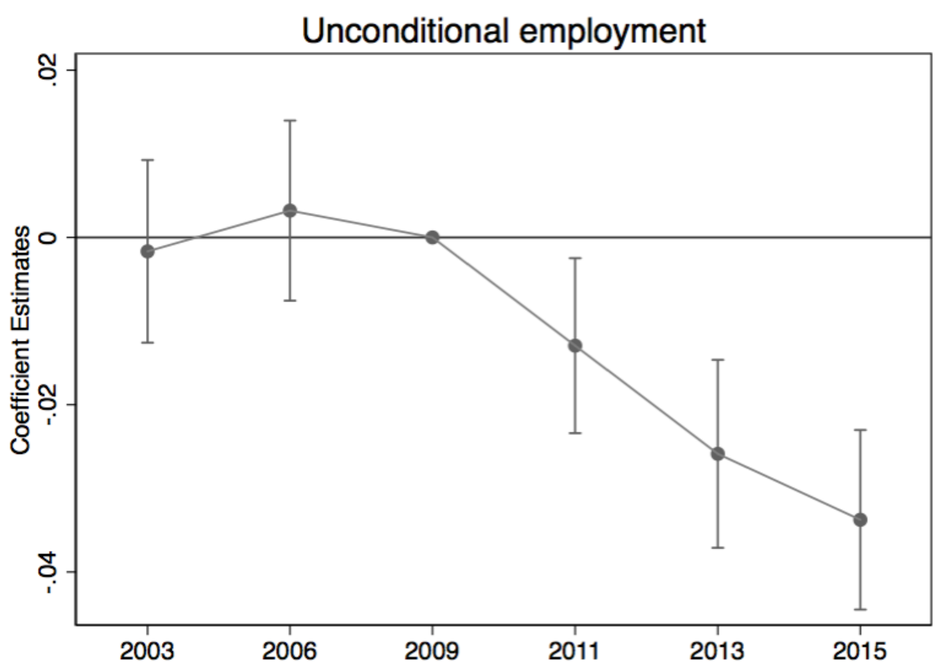
Notes: Point estimates and 95% confidence intervals are plotted. Gender pay gap is estimated conditional on schooling level and experience for each year.

Table 2: Childbearing age women main estimates

	(1) Labor force	(2) Employment	(3) Log Labor Income
Treatment × Post	-0.0307***	-0.0241***	
	(0.00315)	(0.00316)	
Treatment × Post × Female			0.00409
			(0.00850)
Covariates	Yes	Yes	Yes
Year FE	Yes	Yes	Yes
Age dummies	Yes	Yes	Yes
Years of schooling dummies	No	No	Yes
Observations	371440	371440	342898

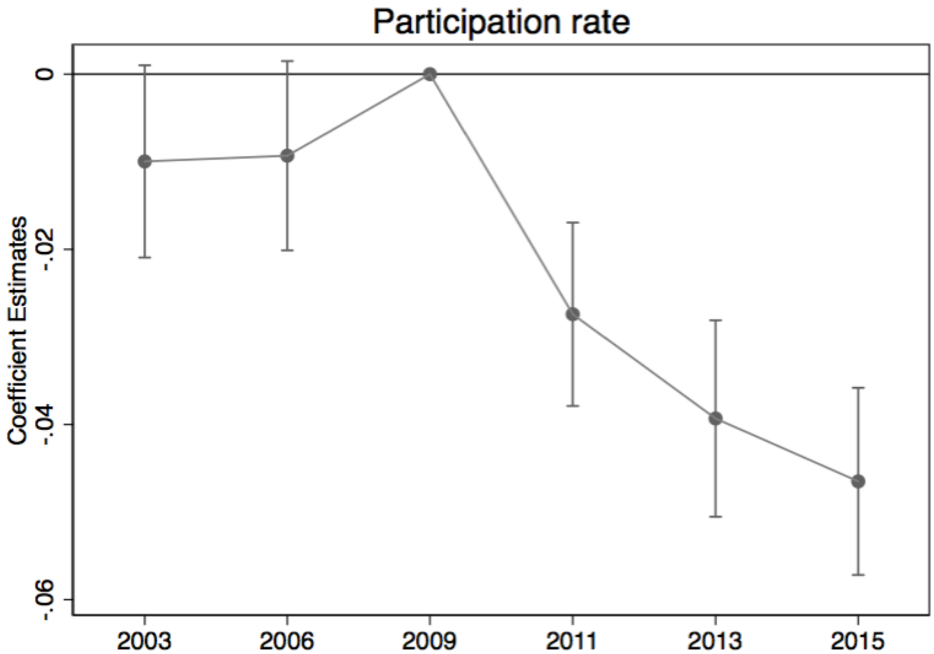
Robust standard errors in parentheses. Column 3 also includes occupation dummies.
* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Figure 8: Event study for unconditional employment



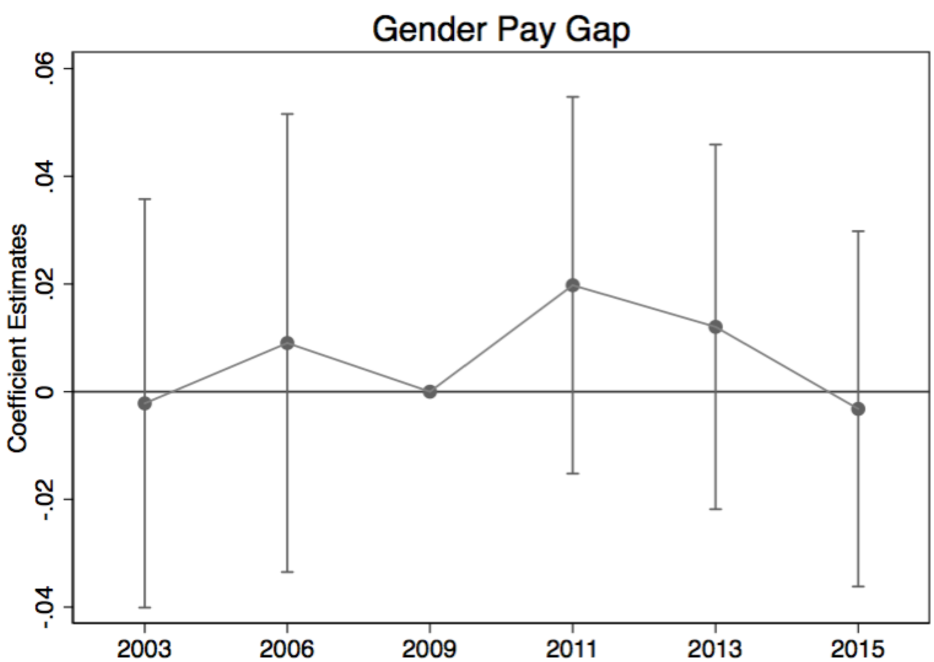
Notes: Point estimates and 95% confidence intervals for each year are plotted. Estimates are relative to 2009, since it is the closest year prior to the implementation of the maternity leave expansion. Treatment group: childbearing age women. Control group: older women (39 to 50 years old).

Figure 9: Event study for labor force participation



Notes: Point estimates and 95% confidence intervals for each year are plotted. Estimates are relative to 2009, since it is the closest year prior to the implementation of the maternity leave expansion. Treatment group: childbearing age women. Control group: older women (39 to 50 years old).

Figure 10: Event study for gender pay gap



Notes: Point estimates and 95% confidence intervals for each year are plotted. Estimates are relative to 2009, since it is the closest year prior to the implementation of the maternity leave expansion. Treatment group: childbearing age women. Control group: older women (39 to 50 years old).

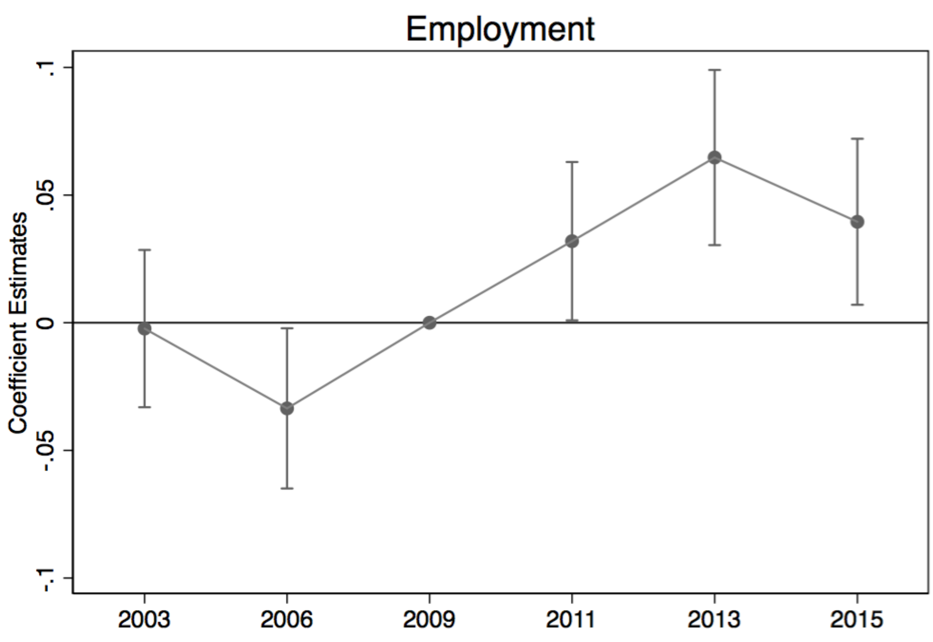
Table 3: Mothers of infants versus mothers of 5- to 10-years-old children main estimates

	(1) Participation rate	(2) Employment	(3) Log labor income
Treated	-0.149***	-0.113***	0.00133
	(0.00679)	(0.00669)	(0.0103)
Treated × Post	0.0580***	0.0554***	-0.0127
	(0.00926)	(0.00919)	(0.0144)
Female			-0.454***
			(0.00821)
Treated× Post × Female			-0.0385
			(0.0289)
Post × Female			-0.0211*
			(0.0104)
Observations	71969	71969	66771

Robust standard errors in parentheses. All columns include year FE, age FE, and controls. Column 3 also includes occupation FE.

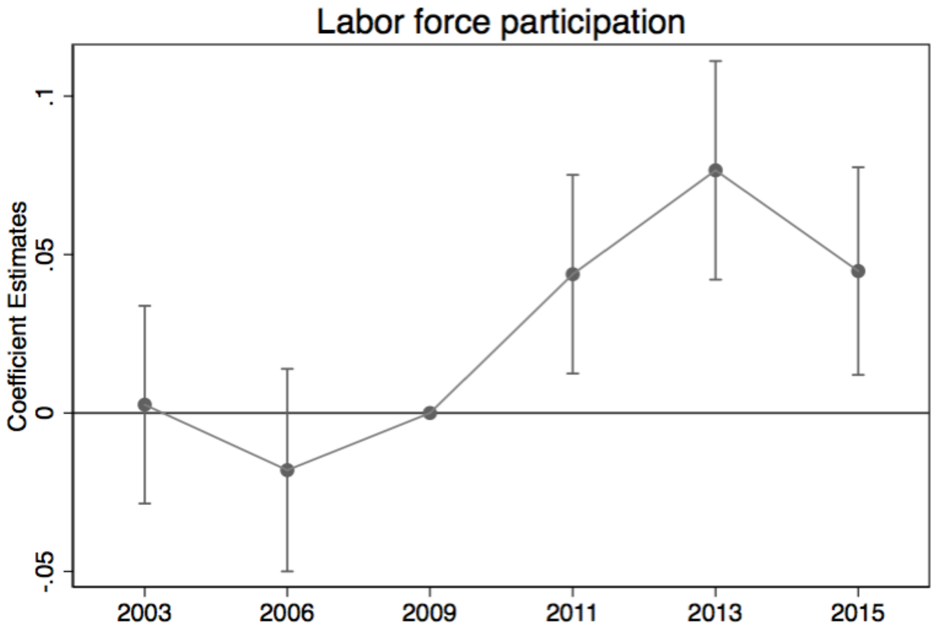
* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Figure 11: Event study for unconditional employment



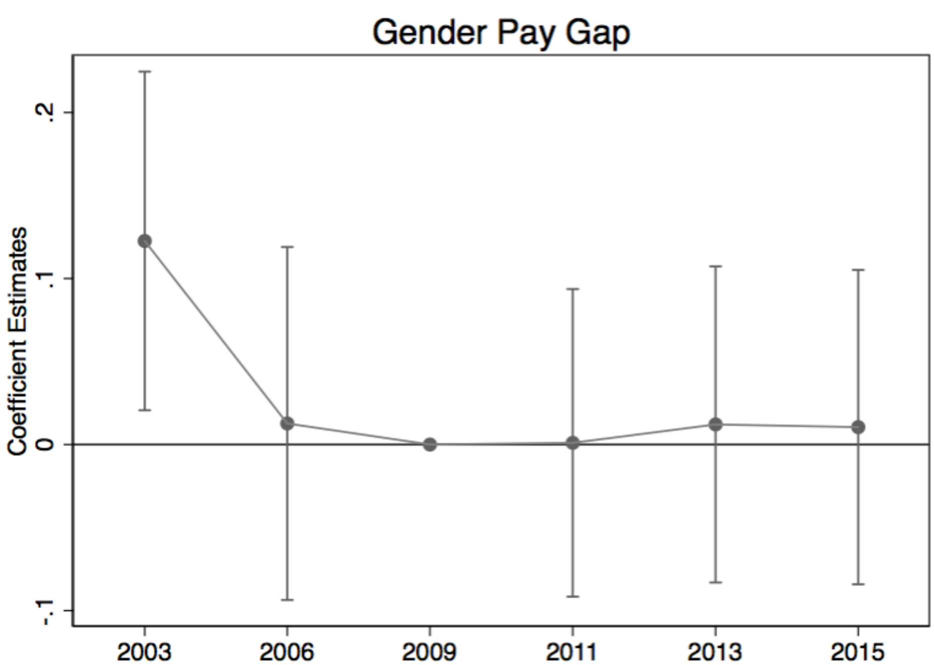
Notes: Point estimates and 95% confidence intervals for each year are plotted. Estimates are relative to 2009, since it is the closest year prior to the implementation of the maternity leave expansion. Treatment group: women having a child younger than 1 year old. Control group: women whose youngest child is between 5 to 10 years old.

Figure 12: Event study for labor force participation



Notes: Point estimates and 95% confidence intervals for each year are plotted. Estimates are relative to 2009, since it is the closest year prior to the implementation of the maternity leave expansion. Treatment group: women having a child younger than 1 year old. Control group: women whose youngest child is between 5 to 10 years old.

Figure 13: Event study for conditional gender pay gap



Notes: Point estimates and 95% confidence intervals for each year are plotted. Estimates are relative to 2009, since it is the closest year prior to the implementation of the maternity leave expansion. Treatment group: women having a child younger than 1 year old. Control group: women whose youngest child is between 5 to 10 years old.

Table 4: Mothers of infants versus childless women main estimates

	(1) Participation rate	(2) Employment	(3) Log labor income
Treated	-0.243***	-0.195***	0.104***
	(0.00646)	(0.00637)	(0.0100)
Treated × Post	0.0996***	0.0894***	0.0235
	(0.00885)	(0.00877)	(0.0138)
Female			-0.203***
			(0.00417)
Treated× Post × Female			-0.0308
			(0.0277)
Post × Female			-0.0527***
			(0.00537)
Observations	220043	220043	206604

Robust standard errors in parentheses. All columns include year FE, age FE, and controls. Column 3 also includes occupation FE.

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Table 5: Mothers of infants versus mothers whose youngest child is between 11 to 18 years old main estimates

	(1) Participation rate	(2) Employment	(3) Log Labor Income
Treated	-0.204***	-0.171***	0.0132
	(0.00796)	(0.00793)	(0.0115)
Treated × Post	0.0622***	0.0566***	0.00765
	(0.00915)	(0.00908)	(0.0145)
Female			-0.430***
			(0.00743)
Treated × Post × Female			-0.0468
			(0.0285)
Post × Female			-0.0209*
			(0.00942)
Observations	83659	83659	73319

Robust standard errors in parentheses. All columns include year FE, age FE, and controls. Column 3 also includes occupation FE.

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Table 6: Mothers whose youngest child is 1 year old versus mothers whose youngest child is between 5 to 10 years old main estimates

	(1) Participation rate	(2) Employment	(3) Log labor income
Treated	-0.114***	-0.100***	0.00745
	(0.00574)	(0.00565)	(0.00845)
Treated × Post	0.0293**	0.0236*	-0.0233
	(0.0101)	(0.0100)	(0.0150)
Female			-0.470***
			(0.00722)
Treated × Post × Female			-0.00653
			(0.0286)
Post × Female			0.0114
			(0.0109)
Observations	72826	72826	67186

Robust standard errors in parentheses. All columns include year FE, age FE, and controls. Column 3 also includes occupation FE.

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Table 7: Mothers whose youngest child is 2 years old versus mothers whose youngest child is between 5 to 10 years old main estimates

	(1) Participation rate	(2) Employment	(3) Log labor income
Treated	-0.0782***	-0.0699***	0.0189*
	(0.00568)	(0.00562)	(0.00803)
Treated × Post	0.0200*	0.0108	-0.0370*
	(0.00990)	(0.00992)	(0.0156)
Female			-0.469***
			(0.00721)
Treated × Post × Female			0.0196
			(0.0282)
Post × Female			0.0114
			(0.0109)
Observations	73331	73331	67793

Robust standard errors in parentheses. All columns include year FE, age FE, and controls. Column 3 also includes occupation FE.
* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Table 8: Mothers whose youngest child is between 5 to 10 years old versus mothers whose youngest child is between 11 to 15 years old main estimates

	(1) Participation rate	(2) Employment	(3) Log labor income
Treated	-0.0492***	-0.0473***	-0.00909
	(0.00432)	(0.00434)	(0.00682)
Treated × Post	0.00795	0.00283	0.0129
	(0.00595)	(0.00602)	(0.0101)
Female			-0.453***
			(0.00939)
Treated × Post × Female			-0.00384
			(0.0165)
Post × Female			-0.0186
			(0.0128)
Observations	99028	99028	86406

Robust standard errors in parentheses. All columns include year FE, age FE, and controls. Column 3 also includes occupation FE.
* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Table 9: Mothers whose youngest child is between 5 to 10 years old versus mothers whose youngest child is between 11 to 18 years old main estimates

	(1) Participation rate	(2) Employment	(3) Log labor income
Treated	-0.0546***	-0.0534***	-0.00474
	(0.00397)	(0.00398)	(0.00626)
Treated× Post	0.00768	0.00292	0.0102
	(0.00536)	(0.00542)	(0.00912)
Female			-0.435***
			(0.00793)
Treated × Post × Female			0.00726
			(0.0149)
Post × Female			-0.0306**
			(0.0106)
Observations	116994	116994	101273

Robust standard errors in parentheses. All columns include year FE, age FE, and controls. Column 3 also includes occupation FE.
* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Table 10: Estimates for mothers of infants by marital status

	Observations	(1) Participation rate	(2) Employment
Single	12247	0.107***	0.0884**
		(0.0311)	(0.0316)
Not single	59722	0.0530***	0.0533***
		(0.00988)	(0.00976)

Robust standard errors in parentheses. Year FE, age FE and controls included in all regressions.
* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Table 11: Alternative specifications estimates: Unconditional employment

	(1)	(2)	(3)	(4)
Treatment × Post	-0.0396***	-0.0396***	-0.0246***	-0.0269***
	(0.00339)	(0.00335)	(0.00322)	(0.00314)
Covariates	No	No	Yes	Yes
Year FE	No	Yes	No	Yes
Age Dummies	No	Yes	No	Yes
Region Dummies	No	No	No	Yes
Observations	371457	371457	371440	371440

Robust standard errors in parentheses

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Table 12: Alternative specifications estimates: Labor force participation

	(1)	(2)	(3)	(4)
Treatment × Post	-0.0443***	-0.0465***	-0.0294***	-0.0340***
	(0.00340)	(0.00336)	(0.00320)	(0.00313)
Covariates	No	No	Yes	Yes
Year FE	No	Yes	No	Yes
Age Dummies	No	Yes	No	Yes
Region Dummies	No	No	No	Yes
Observations	371457	371457	371440	371440

Standard errors in parentheses

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Table 13: Alternative specifications estimates: Gender pay gap

	(1)	(2)	(3)	(4)
Treatment × Post × Female	-0.00249	-0.000349	0.00574	0.00264
	(0.00859)	(0.00866)	(0.00857)	(0.00833)
Year FE	No	Yes	No	Yes
Age Dummies	No	Yes	No	Yes
Schooling Dummies	Yes	No	No	Yes
Region Dummies	No	No	No	Yes
Observations	342898	342898	342898	342898

Robust standard errors in parentheses. All columns include occupation dummies.

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$