Maternity Leave Reform and Female Labor Market Outcomes in Chile*

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March 6, 2018

Abstract

This paper estimates the effects of maternity leave entitlement expansions on maternal labor market outcomes in Chile. We exploit a reform that increased paid leave from 12 to 24 weeks for mothers of children born on July 25, 2011 or later. We implement a Differences-in-Differences (D-D) strategy, using monthly panel matched employer-employee data to assess the effects of the reform on formal employment and wages. We estimate a highly dimensional two-way fixed-effect model that allows us to control for unobserved heterogeneity at the worker-firm level. We find a two percentage points (pp) increase in formal employment for women in the 16-44 age range (childbearing age) in comparison with men in the same age range. On the other hand, females in childbearing age face a decrease in their wages by two percentage points which is consistent with a supply-side driven effect.

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1 Introduction

Paid parental leave policies have been implemented (or enhanced) in many countries to help to balance work and family needs (ILO, 2014). As a consequence, there has been a rebirth of the study of the effect of these type of policies on children outcomes (Albagli and Rau (forthcoming); Dahl et al. (2016); Carneiro et al. (2015); and Danzer and Lavy (2017)). However, evidence on the effects of maternity leave entitlements on female labor market outcomes, especially in middle-income and developing countries, is still scarce.

In this paper, we estimate the effect of a large maternity leave extension on female labor participation and wages in Chile. The reform extended the mandatory paid leave from 12 to 24 weeks for mothers of children born on 25 July 2011 or later. We implement a Differences-in-Differences (D-D) strategy, using a monthly matched employer-employee panel data for nearly 1.5 million workers from 2006 to 2016.[1] The structure and size of our data allow us to both study medium-run effects, and also account for firm heterogeneity in wage setting (Abowd et al., 1999; Card et al., 2016) to obtain a clean effect of the reform on wages.

Using women in childbearing age (16-44) as a treated group and men in the same age range as a control group, we find that the maternity leave extension produced a two percentage points increase in formal female employment. We also find important medium-run effects in employment, with the maternity leave reform having positive impacts on formal employment after five years of the reform. However, when the control group is women in the 45-50 age range there is a decrease in formal employment for women in childbearing age by 1 pp. This is consistent with the results of Bustamante et al. (2015) for the Colombian case.

[1] It corresponds to a 20% sample of the Unemployment Insurance (UI) registry which is a census of the formal private sector.

We also estimate the effect of the leave extension on wages. For this end, we take advantage of the match employer-employee feature of our data that allows us to estimate worker-firm fixed-effect models and therefore account for individual and firm heterogeneity on the wage setting process. Indeed, recent studies have found the importance of firm heterogeneity in wage determination, specifically for women in Portugal (Card et al., 2016) but also for women in Chile (Cruz and Rau, 2017).[2] We find a three percentage points decrease in wages in our preferred specification and it persists in the medium-run.

Our results are consistent with a supply-driven effect of women in childbearing age choosing to work more in the formal sector (than their male counterparts) despite the lower wages because of the paid-leave benefits, as noted in the model proposed by Klerman and Leibowitz (1999). The results are also in line with the estimated effects of maternity leave found for Europe that also show an increase in aggregate female employment, accompanied by a decrease in wages (Ruhm, 1998). Also, for Canada, employment rates for women are shown to increase after a 25 to 50 weeks extension in paid leave (Hanratty and Trzckinski, 2009). However, our results contrast with other studies that find more modest or no effects. Waldfogel (1999) finds no impact on a federal leave policy in the US.[3] Baum (2003) finds positive effects of federal leave policy in the US for employment and negative for wages, but his results are not statistically significant. The main difference with our paper is that the policy analyzed for the US is an unpaid leave policy, while our study focuses on a paid leave extension[4]. The difference in incentives between a paid maternity leave versus an unpaid one might be behind the mentioned differences in the results.

[2]Evidence of the effect of firm heterogeneity in wages is documented for developed countries such as France (Abowd et al., 1999), Germany (Card et al., 2013) and also in developing countries such as Brazil (Alvarez et al., 2018).

[3]Specifically, the Family and Medical Leave Act that guaranteed 12 weeks of unpaid leave.

[4]See Kalb (2018) for a review on unpaid leave policy evaluations

In summary, this paper adds to the existing literature in three ways. First, it provides new evidence on the effect of maternity leave expansions on employment and wages for women using a rich representative panel data from administrative records. Our data allows to control for worker unobserved heterogeneity, and also to look for medium-run effects of the leave reform. Second, we can control for firm heterogeneity on wages since we use matched employer-employee data. This allows us to estimate a highly-dimensional worker-firm fixed effects model *à la* Abowd, Kramarz and Margolis (1999). Finally, we present evidence of maternity leave effect on female labor outcomes for a medium-high income country, which is scarce in the literature.

2 Institutional Background and Data

Maternity leave was first introduced in Chile in 1925. In that period, the leave consisted of 12 mandatory weeks with full income replacement. The leave also guaranteed job protection for one year after completion of the leave. On October 17th, 2011, an extension of 12 weeks was added to the former leave, with the same conditions as the former leave. However, women are allowed to return to work on a part-time arrangement after the 12 mandatory weeks, and also can transfer part of the extended leave to the father.

Thus, the reform entitled mothers who gave birth on or after July 25, 2011, up to 24 weeks of paid leave. We consider women in childbearing age at the time of the reform as the "treatment group" for our study (women between 16 and 44 years old[5]). Given that women use all parental leave in Chile, and as in similar studies (Baker and Milligan, 2008; Baum, 2003; Ruhm, 1998), our main control group will be men of the

[5]According to INE (2016), between 16 and 44 years old is the most accurate estimation of fertile age for females in Chile.

same age as women considered in the treatment group.[6]

The Chilean case proves to be interesting to analyze for two main reasons. First, as noted above, we use new administrative, matched employer-employee data from a developing country, which contributes to the recent knowledge on the subject in a middle-income setting. Second, as explained by Albagli and Rau (forthcoming), the timing of the discussion and the enactment of the reform is unlikely to be endogenous to female workers or mothers. This is because the schedule of the discussion and adoption of the reform: Several legislative proposals were submitted in the past decade, but none were approved. The final legislative process took eight months, with the law approved on October 6, 2011, and finally enacted on October 17, 2011. Thus, firms and female workers could be certain of the content and effect of the reform only on October 2011. This allows to use a differences-in-difference strategy to identify the impact of the reform.

2.1 Data

Our dataset corresponds to a 20% sample of the match employer-employee data from the Unemployment Insurance (UI) registry.[7] By law, the Unemployment Fund Administrator is required to collect and register, on a monthly basis, all contributions to the unemployment insurance individual account for each job and worker. Our data include workers in the formal sector only, which accounts for 70% of the total labor force. This is one of the main features of our data: it's size allows us to build a panel of around 1.5 million workers and 0.58 million firms with a time span of 120 months. The dataset contains worker characteristics such as age, education level, marital status, region, time of affiliation to the unemployment insurance and monthly taxable income. We also have information about employers: number of employers, industry,

[6]In 2017, only 0.2% of fathers use some weeks of parental leave SUSESO (2018).

[7]The is publicly available and can be downloaded at https://www.spensiones.cl/apps/bdp/index.php

and geographic region.

We build a monthly panel from the dataset described above. In Table 1 we present some descriptive statistics (sample means, standard deviations and number of observations of the overall sample). The first and second panel shows information about general characteristics about the individuals in the estimation sample, including our dependent variables formal employment and wages. The average employment rate in the formal sector is 68%, and the average monthly wage is 484,056 CLP (of 2016 pesos), around 716 US dollars. The average age is 33 and around 40% of the sample consist of women. For the sub-sample where we have education data available, we note that around 19% of the sample has completed high-school and 0,7% has completed a degree. While there is probably mis-measurement regarding this information (note that we don't have education information for around 37% of the overall sample), we show estimations using the sub-sample with education information, and the results don't change. The third panel shows the distribution regarding the "treatment status" considered in our study: Our treatment group is defined as the female workers who are in fertile age (less than 44 years old), which represents 33,7% of the sample, and our main control group consists of male workers who are less than 44 years old, which represents 51,8% of the sample. We also use two other control groups which are women between 45 and 50 years old (5,8% of the sample) and finally men between 45 and 50 years old (8,7% of the sample).

3 Empirical Methods

To estimate the effect of the leave reform on females' employment and wages, we first estimate the following lineal probability model for the employment rate:

$$y_{it} = \alpha_i + \theta_t + \beta F_i P_t + \delta F_i + \mathbf{x}'_{it}\phi + \epsilon_{it} \tag{1}$$

where y_{it} is a dummy variable that equals 1 when worker i is employed at period t (month-year), α_i is an individual fixed effect, F_i is a dummy variable that indicates treatment status: it equals to 1 when worker i is a women of fertile age (between 16 and 44 years old) in the observed period and 0 if individual i belongs to the control group (i.e., males between 16 and 44 years old), P_t is a dummy variable that takes the value 1 when the employment status is observed after the reform was implemented (i.e., October 2011) and $\mathbf{x}_{it}$ is a vector of controls such as a quadratic age polynomial, seasonality dummies or time dummies, and educational level interacted with the age polynomial. Finally, ϵ_{it} is an error term, clustered at the individual level. Equation (1) corresponds to a Differences-in-Differences (D-D) model, where β is the parameter of interest.

Exploiting the fact that we have matched employer-employee data, we use the model proposed by Abowd, Kramarz and Margolis (1999) (from now on, AKM) to identify firm fixed effects together with worker fixed effects. This is useful because it allow us to account for time-invariant firm heterogeneity in wage payments. Therefore, we estimate a slightly different model from (1) to address the effect of the leave reform on wages:

$$\ln w_{it} = \alpha_i + \theta_t + \psi_{\mathbf{J}(i,t)} + \gamma F_i P_t + \delta F_i + \mathbf{x}'_{it}\pi + \nu_{it} \tag{2}$$

For Equation (2) we add firm fixed effects $\psi_{\mathbf{J}(i,t)}$, where $j = \mathbf{J}(i,t)$ is the firm where individual i was employed at time t. As noted by AKM, we can only identify firm effects in the "largest connected set" of firms, which correspond to the set of firms connected by the mobility of workers[8] In this case, γ is the D-D estimator for the wage

[8] For example, when a worker moves from firm A to firm B, we say firms A and B are directly connected by mobility. When a worker moves from firm A to firm B, and then another worker moves from firm B to a firm C, then we say firms A and C are indirectly connected by mobility. The largest connected set comprises all firms that are directly or indirectly connected.

equation and the second parameter of interest in this study. Finally, ν_{it} is the error term, also clustered at the individual level.

3.1 Event Study Analysis

We also conduct an event study analysis by generalizing the D-D models in Equations (1) and (2) to take advantage of the panel data. We consider the following equations:

$$y_{it} = \alpha_i + \theta_t + \sum_{t=-5}^{-2} F_i \lambda_t \tilde{\beta}_t + \sum_{t=0}^{5} F_i \lambda_t \tilde{\beta}_t + \mathbf{x}'_{it}\phi + \epsilon_{it} \tag{3}$$

$$\ln(w_{it}) = \alpha_i + \theta_t + \psi_{\mathbf{J}(i,t)} + \sum_{t=-5}^{-2} F_i \lambda_t \tilde{\gamma}_t + \sum_{t=0}^{5} F_i \lambda_t \tilde{\gamma}_t + \mathbf{x}'_{it}\pi + \nu_{it} \tag{4}$$

Equations (3) and (4) are generalizations of Equations (1) and (2) respectively. The omitted dummies in each equation are the interaction between the dummy that indicates treated status F_i and the dummy for being observed one year before the leave reform. Also, we omit the interaction between treatment status and the dummy for being observed in $t = -6$. This in order to avoid under-identification issues as noted by Borusyak and Jaravel (2017). Hence, each $\tilde{\beta}_t, \tilde{\gamma}_t$ can be interpreted as an estimate of the average impact of the leave extension on a given year. Also, each $\tilde{\beta}_s, \tilde{\gamma}_s$ represent D-D estimates for each period before the enactment of the extension reform. Therefore, we should expect these parameters to be near zero under the parallel trends assumption. The advantage of utilizing this unrestricted estimates (and the fact that our data contains a large time span) is that we can look at the dynamics of change in female wages and employment produced by the reform. Also, we can indirectly test for pre-trends in the outcomes of interest and therefore check if the parallel trends assumption is (approximately) satisfied.

4 Results

Table 2 shows the results for four specifications of equation (1) that adds different control variables such as time fixed effects, age, education, and regional dummies. All specifications include worker fixed effects. Column (1) shows the results with worker fixed effects only suggesting a three percentage point (pp) increase in formal employment of women in childbearing age (16-44) compared to men in the same age range. When time fixed effects are included, column (2), the point estimate remains unchanged. Column (3) shows the results with age controls as well and the point estimate reduces to 2.5 pp. Lastly, when educational controls are added in column (4), the point estimates reduces to 2 pp. This is our preferred specification since it controls for a variety of confounding factors. The results are statistically significant at the 1% level.

Regarding the impact of the paid leave extension on female wages, Table 3 shows the results for the estimation of equation (2). Column (1) that includes only worker fixed effects shows a 1,5 percentage points decrease in the logarithm of wages. To account for firm heterogeneity in wage determination, we present estimations of AKM models in Columns (2) to (4) of Table 3, which include firm and worker fixed effects. In this case, we see that the D-D estimator rises in magnitude, finding a 2-4 pp decline in the logarithm of wages. We also present the Adjusted R^2 and the number of worker, firm effects and controls included in the estimation. Again, the estimations are consistent in every specification and statistically significant at the 1% level.

4.1 Dynamics of the Effects

To assess the dynamics of the effects, Figures 1 and 2 show the estimated coefficients of Equations (3) and (4), that correspond to the interaction of the year and treatment status prior to the enactment of the leave reform $(\tilde{\beta}_s, \tilde{\gamma}_s)$ and after the reform $(\tilde{\beta}_t, \tilde{\gamma}_t)$.

In the case of the participation rate for women in childbearing age, we observe from Figure 1 that an increase in employment rate starts two years after the reform has been enacted. We also observe that the employment rate increases every year in larger magnitudes, reaching around a seven percentage points increase five years after the extension was enacted. In years prior to the reform, we can see that coefficients are small and don't reflect a trend in employment, therefore showing that the parallel trends assumption is approximately satisfied for the employment equations we estimate.

When looking at the medium run effects of the leave extension on wages Figure 2 shows that the year of the reform female wages drop nearly two percentage points and then they decrease by smaller amounts reaching four percentage points two years after. The point estimates stabilize in an average decrease of 2.5 pp five years after the reform. Prior to the reform, we observe that most of the coefficients get close to zero, and that they don't represent a trend previous to the reform (i.e., some coefficients are small, with some positive and others negative, to then clearly expose a downward trend after the extension is enacted) therefore showing that for the wage equation the parallel trends assumption is also approximately satisfied.

To go further on the parallel trends assumption Figures 3 and 4 show the evolution of participation and wages for treated and control groups. As can be seen, there trajectories before the reform are parallel in both cases.

4.2 Alternative control groups

Even though most of the literature on the effect of maternity leave on employment use men in the same age range as women in childbearing age as a control group (Baker and Milligan, 2008; Baum, 2003; Ruhm, 1998), in this subsection we assess the effects of the reform on women in childbearing age compared to women in the 45-50 age range (*old women*) and other alternative control groups.

In Tables 4 and 5 we show the results for the preferred specifications of Tables 2 and 3 but with different control groups. While the negative result for wages are robust to different control groups as seen in Table 5, the result in formal employment changes when women in the 45-50 age range are use as a control group instead of men in the 16-44 age range (*young men*). Indeed, column (2) of Table 4 shows that formal employment decreases by 1 percentage for women in childbearing age compared to those in the 45-50 age range. This negative effect in employment when using *old women* as a control group is in line with those of Bustamante et al. (2015) in Colombia. However, while old women could serve as a valid comparison group since they should not be affected by the reform as those in childbearing age, they are in a different stage of their life cycle, facing different labor market prospects that may affect labor market participation. Moreover, different substitution patterns between control groups and young women could be behind the differences in the results. Existing evidence for Italy (Giorgi et al., 2015) and for the US (Acemoglu et al., 2004) shows that young men and women are imperfect substitutes, due to skills and preferences. If women close in age (i.e. women close to 44 years old and women with more than 44 years old) are perfect substitutes, the reform incentives firms to hire more women who are not in childbearing age[9].

Finally, Column (3) repeats the analysis using men in the 45-50 age range (*old men*) finding a reduction of 0.7 percentage points which is smaller than that found in column (2) but significant at the 1% level. Lastly, when old women and young men are used as a control group, column (4), the positive effect dominates and women in childbearing age are 1.7 pp more likely to participate in the formal sector.

In summary, there is a decrease in wages of childbearing age which is robust to different control groups. The effect in formal employment is ambiguous and depends

[9]There is evidence pointing that productivity is not different between close age groups, suggesting perfect substitution between them (Mahlberg et al., 2013)

on the control group used. When *young men* are used as a control group which is our preferred control group, as in several studies (Baker and Milligan, 2008; Baum, 2003; Ruhm, 1998), there is an increase in formal employment. Following the theoretical setting of Klerman and Leibowitz (1999) and the discussion in Ruhm (1998), our results suggest that this movement in supply is driven by the incentives that extension poses for women in childbearing age: 12 weeks more of paid, protected leave makes working in the formal sector more attractive, since the costs of taking care of the child are offset because of the wage, the protection of the formal sector and finally the great reduction in the costs regarding taking care of the (expected) child. There is also empirical evidence for the Australian case supporting the former claim: women are more willing to receive lower wages in exchange of the benefits that paid leave implies (Edwards, 2006). At the same time, a contraction of the demand greater than the shift in supply would explain the employment results when comparing *young women* with different control groups: As noted earlier, if women of close age are perfect substitutes, then formal employment for *young women* would decrease compared with *old women* after the reform. This is because firms could perceive the mandatory leave policy as an extra cost of hiring women in childbearing age, leading them to hire more women who are slightly older, despite the entry of women willing to work for lower wages.

5 Conclusions

The effects of maternity leave expansions have been extensively studied in the last decades. Most of the evidence comes from developed countries and the lack of matched employer-employee data at the time these reforms were implemented, makes more difficult to control for unobserved heterogeneity at the firm-worker level.

This paper estimates the effects of maternity leave entitlement expansions on maternal labor market outcomes in Chile. We exploit a reform that increased paid leave from 12 to 24 weeks for mothers of children born on July 25, 2011 or later. We implement a Differences-in-Differences (D-D) strategy, using monthly panel matched employer-employee data to assess the effects of the reform on formal employment and wages. We estimate a highly dimensional worker-firm fixed-effect model finding a two percentage points increase in formal employment for women in the 16-44 age range (childbearing age) in comparison with men in the same age range. On the other hand, females in childbearing age face a decrease in wages by three percentage points which is consistent with a supply-side driven effect.

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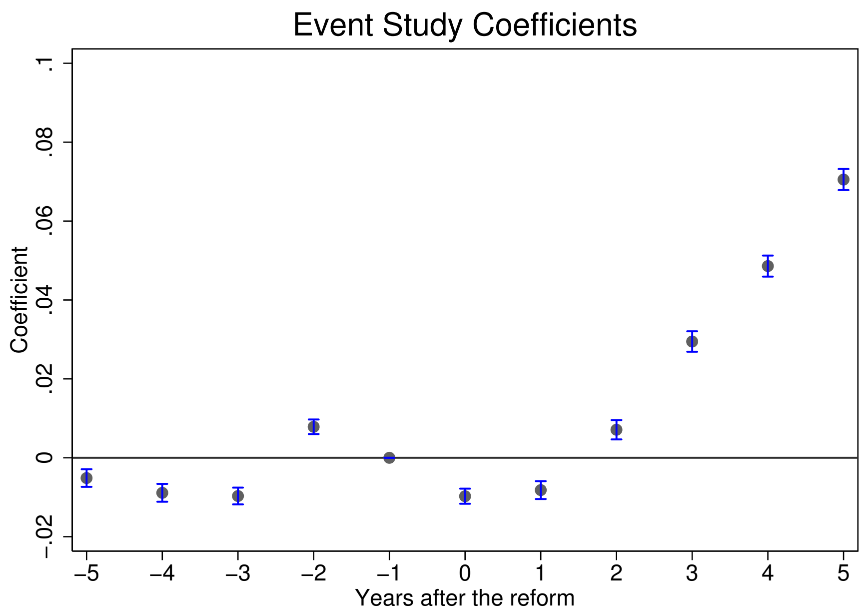

Figure 1: Medium-Run effects of Maternity Leave Extension in Female Participation in the formal sector: Estimated coefficients of year and treatment status interaction (and confidence intervals) are shown.

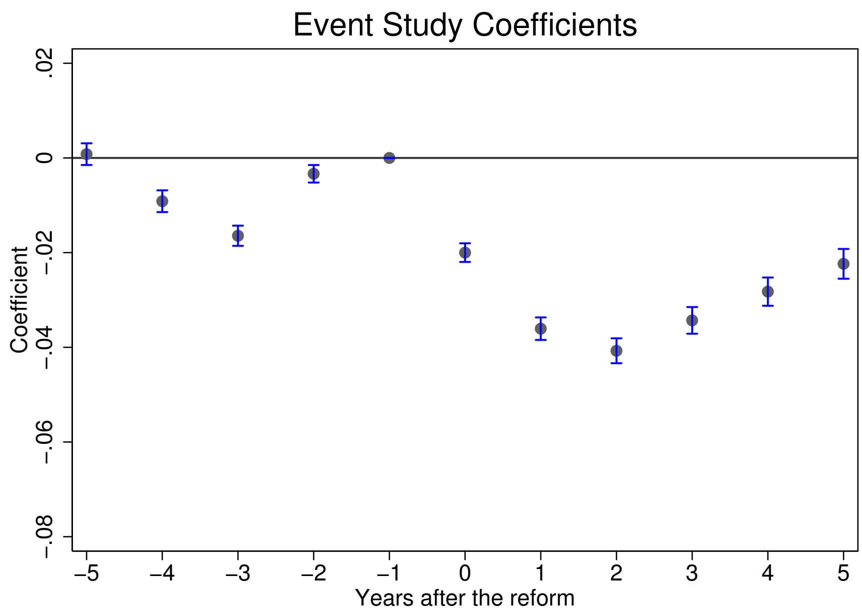

Figure 2: Medium-Run effects of Maternity Leave Extension in Female Wages: Estimated coefficients of year and treatment status interaction (and confidence intervals) are shown.

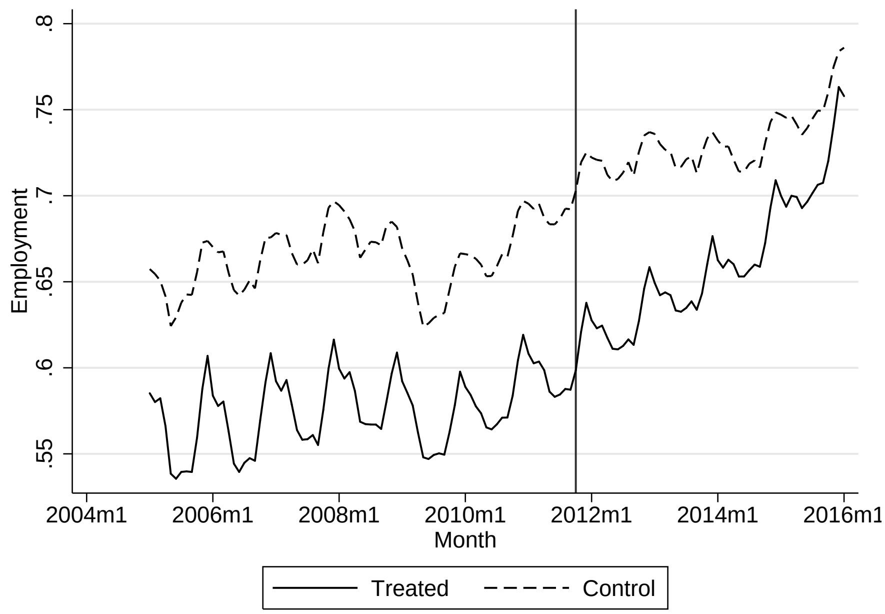

Figure 3: Formal Sector Employment of Treated (women in childbearing age) and Control (men of less than 45 years old) Groups, monthly series

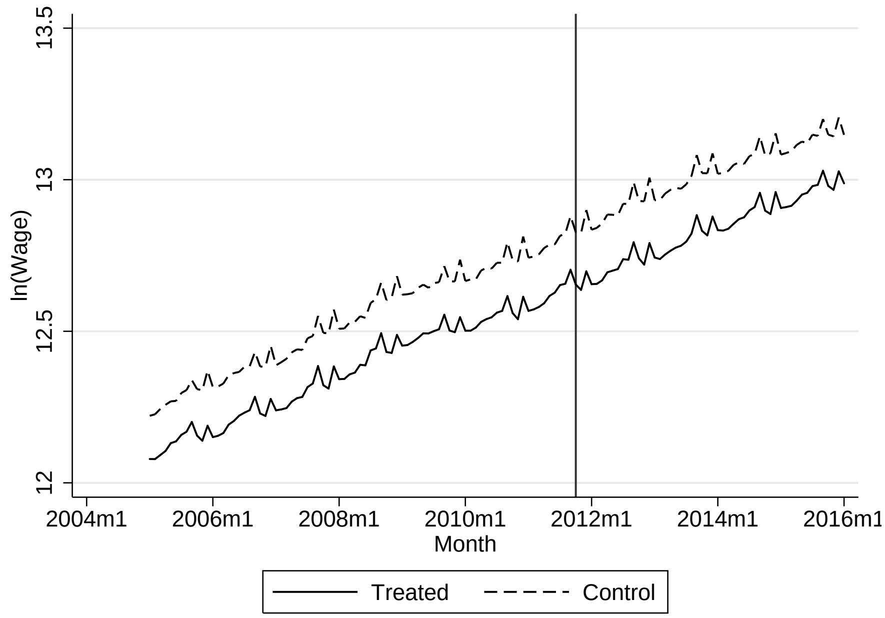

Figure 4: Wages of Treated (women in childbearing age) and Control (men of less than 45 years old) Groups, monthly series

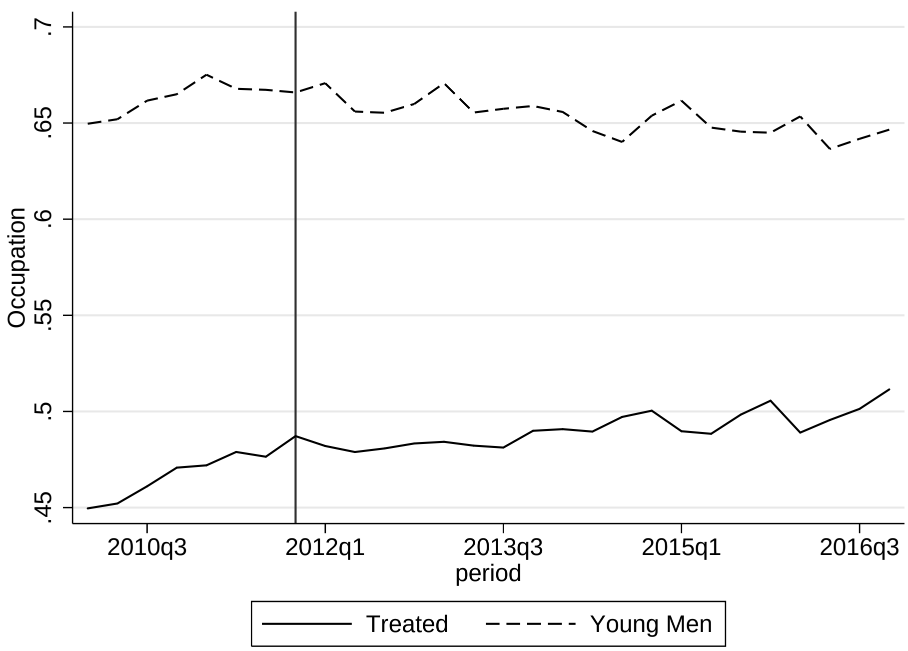

Figure 5: Employment of Treated and Control Groups

Figure 6: Employment of Treated Women and Older Women

Table 1: Descriptive Statistics

	(1) Mean/Share	(2) Std. Dev.	(3) N
Is Working (Formal Sector, Monthly)	0.683	0.465	132,854,426
Wage of Workers	484056	488160	90,682,963
ln(Wages)	12.727	0.823	90,682,963
Age	33	8.846	132,854,426
Women	0.395	0.489	132,854,426
Men	0.610	0.489	132,854,426
Education Controls:			
Completed High-School	0.189	0.391	97,970,724
Completed Degree	0.007	0.084	97,970,724
Treatment Status:			
Women in Fertile Age ($\leq$ 44 years old)	0.337	0.473	132,854,426
Young Men ($\leq$ 44 years old)	0.518	0.499	132,854,426
Older Women (between 45 and 50 years old)	0.058	0.233	132,854,426
Older Men (between 45 and 50 years old)	0.087	0.282	132,854,426

Table 2: Effects of Leave Extension on Employment

	Outcome is Probability of Being Employed			
	(1)	(2)	(3)	(4)
Maternity Leave D-D	0.029***	0.029***	0.025***	0.020***
	(0.001)	(0.001)	(0.001)	(0.001)
Worker FE	Yes	Yes	Yes	Yes
Time FE	No	Yes	Yes	Yes
Age Controls	No	No	Yes	Yes
Education Controls	No	No	No	Yes
Observations	113,565,784	113,565,784	113,565,784	83,907,970

Clustered standard errors at the individual level in parenthesis. * $p<0.1$, ** $p<0.05$, *** $p<0.01$.

Table 3: Effects of Leave Extension on Wages

	Outcome is Logarithm of Monthly Wage			
	(1)	(2)	(3)	(4)
Maternity Leave D-D	-0.003***	-0.005***	-0.010***	-0.023***
	(0.001)	(0.001)	(0.001)	(0.001)
Firm FE (AKM)	Yes	Yes	Yes	Yes
Worker FE	No	Yes	Yes	Yes
Age Controls	No	No	Yes	Yes
Education Controls	No	No	No	Yes
Adjusted R^2	0.64	0.83	0.83	0.80
Root-MSE	0.53	0.36	0.36	0.37
Firm Degrees of Freedom	506,236	465,176	465,176	395,742
Individual Degrees of Freedom		1,396,066	1,396,066	848,326
Covariates Degrees of Freedom		12	13	48
Observations	76,527,007	76,434,883	76,434,883	57,641,646

Clustered (at worker level) standard errors at the individual level in parenthesis. * $p<0.1$, ** $p<0.05$, *** $p<0.01$. All equations control for a linear time trend and seasonality dummies. In AKM (Firm and Worker Fixed Effect) specifications, singletons are dropped.

Table 4: Effect of Leave Extension on Employment: Different Control Groups

	(1) Young Men	(2) Old Women	(3) Old Men	(4) Old Women and Young Men
Maternity Leave D-D	0.020***	-0.011***	-0.007***	0.017***
	(0.001)	(0.001)	(0.001)	(0.001)
Observations	83,907,970	35,723,850	26,634,327	89,181,068

Clustered standard errors at the individual level in parenthesis. * p<0.1, ** p<0.05, *** p<0.01. Column (1) uses Men in the same age as treated women as the control group, is equivalent to the Column (4) of Table 1. Column (2) uses Women not in childbearing age as the control group. Column (3) uses Men "not in childbearing age" as a control group. Column (4) uses Men with the same age as treated women and Women not in childbearing age as a control group.

Table 5: Effect of Leave Extension on Wages: Different Control Groups

	(1) Young Men	(2) Old Women	(3) Old Men	(4) Old Women and Young Men
Maternity Leave D-D	-0.039***	-0.012***	-0.018***	-0.039***
	(0.001)	(0.001)	(0.001)	(0.001)
Observations	57,713,842	24,994,529	30,275,699	63,095,487

Clustered standard errors at the individual level in parenthesis. * $p<0.1$, ** $p<0.05$, *** $p<0.01$. Column (1) uses Men in the same age as treated women as the control group, is equivalent to the last Column of Table 1. Column (2) uses Women not in childbearing age as the control group. Column (3) uses Men "not in childbearing age" as a control group. Column (4) uses Men with the same age as treated women and Women not in childbearing age as a control group.

Appendix

Table 6: Summary of AKM Estimation

	(1)
Std Dev of ln(Wages)	0.823
Worker-Year Obs	76,434,883
Summary of Estimated Parameters:	
Number of Worker Effects	1,396,066
Number of Firm Effects	465,176
RMSE of Estimation	0.353
Adjusted R^2	0.815
Std Dev of Worker Effects	0.494
Std Dev of Firm Effects	0.358
Std Dev of $\mathbf{x}'\hat{\phi}$	0.365
Correlation between Worker and Firm Effects	0.251
Match Effects Model:	
RMSE of Estimation	0.289
Adjusted R^2	0.877
Decomposition of Inequality:	
Share of Variance due to:	
Worker Effects:	36,1 %
Firm Effects:	18,9 %
Corr. of Worker and Firm Effs:	6,5%
Share of $\mathbf{x}'\hat{\phi}$	20,8 %
Share of Residuals	17,7 %