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Do Generous Parental Leave Policies Help Top Female Earners?

Abstract

Generous government-mandated parental leave is generally viewed as an effective policy to support women's careers around childbirth. But does it help women to reach top positions in the upper pay echelon of their firms? Using longitudinal employer-employee matched data for the entire Norwegian population, we address this question exploiting a series of reforms that expanded paid leave from 30 weeks in 1989 to 52 weeks in 1993. The representation of women in top positions has only moderately increased over time, and career profiles of female top earners within firms are significantly different from those of their male counterparts. The reforms did not affect, and possibly decreased, the probability for women to be at the top over their life cycle. We discuss some implications of this result to put into perspective the design of new family-friendly policy interventions.

JEL-Codes: J180, J210, J220, J240, M140.

Keywords: top earners, parental leave, women, regression discontinuity.

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I. Introduction

Over the past century, the barriers to women's educational and labour market opportunities have been substantially reduced in most advanced economies. Gender gaps that favour women have even opened up in the case of education, from primary school performance to university graduation, with women who are typically more engaged, spend more time reading, study harder, and achieve more than their male counterparts (Goldin, Katz, and Kuzmienko, 2006; OECD, 2015; Francesconi and Parey, 2018). Women, however, continue to have lower rates of labour force participation and earn less than men, despite the fact that the wage gap has narrowed considerably since the 1960s (Bertrand, 2010; Goldin, 2014; Olivetti and Petrongolo, 2016; Blau and Kahn, 2017; Bertrand and Duflo, 2017). Strikingly, gender differences in wages and earnings seem to be particularly pronounced among top earners (Albrecht, Björklund, and Vroman, 2003; Arulampalam, Booth, and Bryan, 2007; Bertrand, Goldin, and Katz, 2010; Albrecht et al., 2018), and women tend to be underrepresented in top positions (Miller, 2018). In this study, which focuses on Norway, we first highlight the extent of women's representation in top positions and then ask whether more generous leave mandates, in terms of length of leave and wage replacement, have helped women to reach or maintain top earnings positions.

Highly paid women are typically highly educated, highly skilled workers who have heavily invested in their human capital accumulation and are strongly attached to the labour market and their careers (Bihagen, Nermo and Stern, 2014). Promotion tracks and career progression are generally very steep for workers in these groups, where missing a set window of opportunity for advancement may close any chance of catching up (Lazear and Rosen, 1990; Demougin and Siow, 1994). Focusing on such workers, therefore, is likely to allow us to assess more directly whether parental leave policies have a strong positive impact on careers and wages, since we expect labour market participation of high paid women to be highly responsive to work interruptions due to motherhood (see, for instance, Mincer and Polachek, 1974; Becker, 1975; Francesconi, 2002; Adda, Dustmann, and Stevens, 2017).

There is overwhelming evidence that women are underrepresented in highly paid jobs and top corporate positions worldwide (Miller, 2018). Besides career interruptions due to child bearing and rearing (Gronau, 1973; Weiss and Gronau, 1981; Mincer and Ofek, 1982; Becker, 1991), other explanations have been analysed. These include lower chances of internal promotion (Kunze and Miller, 2017; although Francesconi, 2001 and Booth, Francesconi, and Frank, 2003 provide mixed evidence on this explanation for the UK), statistical discrimination (Correll, Benard, and Paik, 2007; Bjerk, 2008; Gayle and Golan, 2012), and stereotyping (Beaman et al. 2009; Bertrand, 2010). Bertrand, Goldin, and Katz (2010) find that female professionals holding a master's degree in business administration (MBA) from a top US business school face a substantial earnings disadvantage compared to their male counterparts ten years after MBA completion, even though male and female MBAs have nearly identical earnings at the outset of their careers. Two key factors that account for this large gender gap are greater career discontinuities and shorter work hours for female MBAs, both of which are strongly associated with motherhood. In contrast, Azmat and Ferrer (2017) find no (negative) effects of having children on performance, hours billed, and client revenue among associate lawyers in the US.

The experience of female corporate leaders and top professionals in the US, however, may be hard to generalise, given they face relatively short parental leave arrangements. But looking at the impact of gender-neutral tenure clock stopping policies introduced in the top-50 US departments of economics (in which untenured assistant professors are allowed to stop their tenure clock for one year after childbirth or adoption), Antecol, Bedard, and Stearns (2018) find these policies decrease the probability that a female assistant professor gets tenure where she was initially hired by 22 percentage points while male tenure rises by 19 percentage points. It seems that, after the policies were implemented, men publish more in prestigious (top-5) journals, while women do not; and women earn tenure somewhere else. Although the objective of these policies might have been to increase family-friendliness and level the playing field for women, these results suggest they are not successful. Moreover, gender pay disadvantages have also been found among Swedish women (Angelov, Johansson, and Lindahl, 2016) and Norwegian female top earners (Bütikofer, Jensen, and Salvanes, 2018), who typically enjoy longer and more generous maternity leaves than their US counterparts.

The empirical evidence on the efficacy of maternity leave policies to safeguard all women's (not only those in highly paid) careers around childbirth is mixed. In the case of relatively short leaves (of six weeks or less), as it is the case in Canada and the US, there is evidence in support of an increased probability for all women to return to the same employer and increases in work hours after a spell of maternity leave (e.g., Baker and Milligan, 2008; Rossin-Slater, Ruhm, and Waldfogel, 2013; Baum and Ruhm, 2016). The experience of women in European countries with longer leaves, however, leads to different conclusions. For instance, exploiting exogenous variation in the length of parental leave generated by policy changes in the German labour market, Ejrnæs and Kunze (2013) find that leave durations of up to 36 months led to detrimental effects on employment and wages for women. Looking at the same changes in Germany, Schönberg and Ludsteck (2014) confirm that expansions in leave coverage reduce mothers' post-birth employment rates in the short run, although the negative longer-run effects seem to be more modest. Lalive et al. (2014) document that extending both cash benefits and job protection above 12 months in parental leave arrangements in Austria induces a significant postponement in mother's decision of returning to work, regardless of whether this means returning to the same pre-birth employer or delaying the return to a new job. Focusing on the pre-1993 policy reforms in Norway, Dahl et al. (2016) find that the expansions in government-funded maternity leave from 18 to 35 weeks had little effect on a wide variety of outcomes, including parental earnings and participation in the labour market in the short or long run, completed fertility, marriage, and divorce. For recent comprehensive surveys, see Kunze (2016), Olivetti and Petrongolo (2017), and Rossin-Slater (2018).

All of the maternity leave studies discussed above focus primarily on all women. Only few consider highly educated or highly skilled (but not necessarily highly paid) mothers. Among them, Cipollone, Patacchini, and Vallanti (2014) show that female labour force participation of medium- and highly-educated women across 15 EU countries is more responsive to family-oriented policies – as measured by a synthetic index encompassing parental leave, family subsidies, and elderly subsidies – than participation of less-educated women. This study does not look at earnings. Ejrnæs and Kunze (2013) find substantial parental-leave-induced wage losses among highly skilled women in Germany. Using aggregate country-level data from about 30 OECD economies, Olivetti and Petrongolo (2017) find evidence that highly educated women see their relative earnings fall as a results of longer maternity leave entitlements.

Our main contribution is to analyse the impact of maternity leave expansions on female earnings, focusing on women in top earnings positions within their organisations. In particular, we ask whether the expansionary parental leave policies observed in Norway in the early 1990s have helped women in the short- and long-term to reach high paying positions in their firms. We use a new measure of top jobs defined on the basis of being in the top decile (or top quartile) of the *within-firm* earnings distribution. This measure could underrepresent women who may be in high positions if their firms systematically underpay them relative to men who are in the same ranks or even in lower positions. Put differently, it is possible that some women in high-ranked positions in some firms earn less than men without such a high rank in their organisations. But to the extent that pay reflects internal career success sufficiently well, our measure is likely to capture the actual economic value that firms assign to their employees, men and women alike (e.g., Baker, Gibbs, and Holmstrom, 1994; Bertrand and Hallock, 2001; Lazear and Oyer, 2004). In future work, we plan to assess the correlation of this measure based on within-firm relative earnings with other measures based on job title or occupation.

Using register data that cover the universe of the working population, we analyse Norway, a country characterised by high female labour force participation, a generous social insurance system, and a greater and more generous maternal leave provision than that granted in many advanced economies, including the United States, Canada, Japan, Australia, and the United Kingdom. We focus on the 1993 parental leave reform, which came at the end of a series of policy changes that steadily expanded paid leave from 30 weeks (in 1989) to 52 weeks with 80% wage replacement or, with full wage replacement, from 20 weeks in 1987 to 42 weeks. We stress two important descriptive results and one new substantive finding. First, we find that the 1993 reform led to an increase in average duration of leave for women, even though already before the reform virtually all mothers took up their maternity leave and many opted for the maximum length. Second, women are largely underrepresented at the top of their firms' pay distribution, with less than 2% being in the top decile from the mid-1980s to the mid-2010s, thus accounting on average for only one-fifth of workers at the top. Hence, also for Norway, one of the most gender equal countries in the world, we find only a modest progress in the percentage women at the top. Third, our main finding from the empirical regression analysis is that the 1993 expansion in paid maternity leave did not affect, and possibly decreased, the probability for mothers to be or move to the top earnings decile and quartile within their firms. This impact emerges soon after childbirth and remains stable over the next 20 years.

After we describe the relevant institutional environment in Norway in Section II and the data in Section III, Section IV elaborates some implications from a simple economic model and presents the econometric methods used in the analysis. We show our main results in Section V and use them to provide a few pointers for public policy discussions in Section VI.

II. Parental Leave in Norway Before 1993 and After

The government-mandated parental leave system in Norway was introduced in 1977. For an overview of the parental leave reforms in Norway, see Table 1. Working women who gave birth between July 1977 and April 1987 were eligible to 18 weeks of paid maternity leave. The law gave them also the right to job protection, that is, to return to the same employer at the same position they had before childbirth. In other words, they could not be laid off over the entire duration of their leave. Since 1978 parental leave could be shared by the parents, but there was no explicit paternal quota.

The 1977 statutory default allowed women to have three weeks of leave before the expected date of childbirth, and 6 weeks thereafter. The remaining 9 weeks could be shared by the father and the mother. The 18 weeks of leave were fully paid. In order to qualify, mothers had to work six out of the 10 months before the birth. Fathers' eligibility required both they and their partners worked at least six out of the 10 months before the birth. Over the same period, income for both mothers and fathers had to exceed the basic income level. Parents obtained compensation for income loss equivalent to average earnings for the three months before the start of the protected leave period, with a cap of six times the government-set basic income up to which full wage replacement was paid. In our sample, only 9-10% of women earn more than the cap at any point in time. However, employers in the public sector topped up the benefits so that their female employees had 100% wage replacement, while employers in the private sector could do so voluntarily. Although there are no reliable data on the extent of supplementary pay, anecdotal evidence suggests that this was common practice. Compensation was – and still is – paid out of tax-payer contributions. Payments to the mother were based on her pre-leave earnings, whereas father's payments were reduced proportionally if the mother did not work full-time before birth.

Since May 1987, the parental leave duration has been gradually expanded by at least 2 weeks almost every year, up to the larger reform which took place in April 1993. From April 1989 onwards, parents can choose either 100% compensation over a shorter leave window or a longer leave with a wage replacement of 80%. By April 1993, the leave duration was extended to 42 weeks at 100% replacement rate or 52 weeks at 80% replacement. The 1993 reform is the final reform in this sequence of expansionary reforms and is the main focus of our analysis. (We have also performed a number of tests on the reforms from 1987 to 1992, but for space limitations, we do not report results on them. They are broadly consistent with what we show for the 1993 reform.)

Figure 1 displays the duration (in days) of the maternity leave provision of the 1993 reform for the 6-month window around April, when the reform was introduced. Childbirth month and year are normalized to 0 for that month. We observe a significant and sharp jump up around the April cutoff date just short of the 3 additional weeks offered to mothers by the 1993 reform. This provides clear evidence that the 1993 reform had a major impact on women' behaviour, with virtually all of them opting for a full leave.

III. Data

Sample and Outcomes – Our analysis uses employer-employee matched panel data, extracted from multiple registers on the entire population of full- and part-time employees and firms in Norway, following their complete employment and earnings histories from 1967 to 2013. To analyse females in top positions *within* firms, we focus on annual earnings reported in the tax registers. We cannot pin down individuals in supervisory and executive boards unless they are officially in paid employment. This is because, in general, directors on boards are not employees of the firm. Earnings are deflated using the Norwegian consumer price index (base year is 1998 = 100), and individuals with very low earnings (i.e., earnings below the annually adjusted basic income according to the Norwegian social security system) are excluded. Including them does not affect our results. Our main outcome measure is given by an indicator variable that takes value 1 if woman i is in the top decile of the earnings distribution in her *own* firm, and 0 otherwise. We label this variable $y_i^{(10)}$. We also consider, and present results for $y_i^{(25)}$, which identifies women in the top quartile of the earnings distribution within their firms. This measure captures a broader set of female workers who may be in high-ranked positions but could earn less than lower-ranked male colleagues in their organisations. Both measures can be calculated consistently across all years available in the registers.

Figure 2 presents the averages for the proportion of women among the top 10%, $y^{(10)}$, and top 25%, $y^{(25)}$. For comparison, we also plot female employment as a fraction of overall (male and female) employment over the sample period. The female proportions in the workforce in Norway, which are among the highest in OECD countries, have remained stable around 48% with respect to male employment from 1987 to 2001 and grown by about 1 percentage point ever since. Women at the top of the pay distribution in each firm, however, are severely underrepresented. For most of the period, there have been less than 2% of women in the top decile and less than 8% in the top quartile, thus accounting for at most one-fifth and one-third of the workers in each part of the distribution, respectively. Despite this lower representation at the top within firms, the trends have been increasing, indicating a steady – albeit slow – gender convergence. Looking at the top quartile, where women in middle management positions are more likely to be, we observe a relatively stronger increase than at the very top. This result is in line with the findings reported in Fortin, Bell and Böhm (2017) and Atkinson, Casarico, and Voitchovsky (2018) on top incomes. For reference, in 2020, only 7% of the CEOs in the largest firms in Norway are women.

Other Variables – In the analysis, we use other individual characteristics, such as gender, age, marital status, and years of education. All variables are assembled from 1987 to 2013 and from various Norwegian administrative registers such as the education registry, the birth registry, the population and employment registries. Work experience is defined as the accumulated number of years with earnings above the relevant annual basic income. The multigenerational birth registry contains information from birth certificates, where both mother and father are reported:

from this registry, we identify the complete record of timing and number of births up to 2013. From the welfare registers, we obtain information on whether and when a woman takes her maternal leave. Parental leave duration is constructed as the total number of days spent in parental leave in relation to each childbirth event. Moreover, since we know the exact number of children, parity, and calendar months of childbirths, we can link each parental leave episode to the exact parity of each child.

Firms can be followed over time owing to a unique plant identification number. The list of plant identifiers, however, has a break in 1995. We thus use the employer-employee matched data and check if workers in the same plant are the same from year to year. Based on this careful exercise, we re-define a new firm identifier that is consistent over the whole sample period. Hethey-Maier and Schmieder (2013) provide more information on the programming routine we use. Years of tenure are measured as the accumulated number of years with earnings above the annual basic income within the same plant identifier. Based on the plant identifiers, we can establish the detailed industry in which each firm in the data operates at the 3-digit level NACE classification.

Descriptives – To describe careers toward the top echelons and gender differences, we report the means for selected variables in Table 2 for the birth cohorts born between 1952 and 1967 in our sample. We compare women in the top decile of their firms' earnings distribution to all other women (i.e., those in the bottom 90%) as well as to men in the top decile. We present one set of figures for 1992, the year just before the 1993 reform, and another for 2013, the last year in our sample. These years could be seen as representing the early and the more senior career stages for the women in this analysis. In 1992, compared to the bottom 90% of female employees, women in the top decile are more educated, older when they become mothers, and less likely to be married, and have fewer children and more work experience. Although they are slightly more likely to take maternity leave (albeit not significantly so), they tend to go back to their jobs 3-4 weeks earlier than their lower-pay counterparts. In 2013, when all women have become older, some of these differences persist, while others (e.g., in age, marital status, and number of children) go in the opposite direction or disappear (as in the case of work experience).

It is also interesting to compare the profiles women have at the top to men at the top. We find that women are younger and more educated than men. However, they are also, less likely to be married, younger when they become parents, have fewer children, less work experience, and shorter firm tenure. Along all characteristics and in both years, these gender differences are always statistically significant, even though they tend to be smaller at the end of the observation period.

IV. Theoretical Considerations and Empirical Strategy

Conceptual Framework – Before moving to the empirical methodology, we provide an intuition of the predictions on employment and wages from expanding family leave, predictions which will guide the interpretation of our results. Building on a standard static labour supply model, we take the setup developed by Klerman and Leibowitz (1997), which explores the labour supply effects of government-mandated maternity leaves. Goldin, Kerr, and Olivetti (2020) formulate a two-period version of the model and look at firm-provided paid parental leave. With no mandate, employers may voluntarily offer an unpaid (or paid) maternity leave. This could be due to private incentives to preserve good matches and exploit greater returns to job-specific human capital. Females choose between this leave and severing the employment relationship to be at home with their child for a longer period and possibly joining another firm at a later date. The cost of quitting is the difference between the current wage and the alternative wage. Females make an optimal choice given a reservation wage that declines with each month after giving birth.

A government leave mandate, such as the one we focus on in the analysis, which exceeds the employer's voluntary offer, will lead some females who previously would have quit their jobs to remain employed and take the mandated leave. Also, some females who had previously taken the shorter leave offered by the employer will now take the additional weeks allowed under the mandate. Therefore, the mandate will reduce the number of women quitting their jobs pre-birth to spend time at home, and increase the number of women who are employed and on leave over the birth event. The model then delivers a positive employment effect, although this is not a margin of behaviour we will explore.

The model instead has no definitive prediction for the average amount of time women are at home with their child, as some mothers may take longer leaves but others shorter leaves, depending for instance on preferences and child idiosyncratic shocks. This will have heterogeneous impacts on women's human capital depreciation. If women in high-pay posts depreciate their human capital more quickly, we would expect them to return to work earlier than their low-pay counterparts. But, as documented in Figure 1, this is not the case in Norway on average, with most mothers using up their full leave entitlements. Human capital depreciation will affects wages, with longer work interruptions leading to a greater decline in earnings. If all mothers use full leaves and highly paid women depreciate their human capital more than low-pay women, their earnings are likely to grow less. Similarly, if high wage jobs are characterised by a pay structure that is nonlinear in hours worked (i.e., working, say, 60 hours per week leads to a pay that is more than twice as much as the pay enjoyed if working 30 hours per week), longer work interruptions or more frequent work absences due to early child care could result in lower earnings. If human capital depreciation or nonlinear pay structures are at work, therefore, we expect to see an increase (or no change at best) in female underrepresentation at the top of the pay distribution within their firms.

Other considerations, besides human capital depreciation and nonlinear pay structures, may play an important role for wage growth over the life cycle as a result of the improvement in mandated leave. With job security no longer a concern, mothers may change jobs to pursue careers with larger wage gains; but job changers could disproportionately be women who anticipate less wage progression in their current firm. The net impact of job turnover on top earnings representation of mothers is therefore ambiguous. Although in our empirical analysis we will not focus on job mobility, this is an interesting area for future research.

Moreover, employers may respond to the leave mandate in altering their hiring process away from women with higher chances of having a birth, or reducing their investment in the career of women who they believe are more likely to leave to raise families (Thomas, 2017; Corekcioglu, Francesconi, and Kunze, in preparation (a)). This may have long term career effects.

Empirical Design – Comparing the career outcomes of working women who have taken shorter and longer maternity leaves is problematic because the decision of maternity leave duration is not random. For instance, women who decide to take longer parental leave may be more family oriented or have less ambitious career goals. To shed light into the effects of parental leave duration on the labour market performance of women at the top of the earnings distribution, we use a regression discontinuity (RD) design method, in which women differ only in their eligibility to longer maternity leaves by a small margin (Lee and Lemieux, 2010).

Specifically, we compare women who gave birth just after April 1, 1993, who are eligible to 42 weeks of leave at 100% salary replacement rate (or 52 weeks at 80% replacement), to women who gave birth just before, who were entitled to only 35 weeks at 100% (or 44.4 weeks at 80%) wage replacement. The idea is that there is random variation in career outcomes among women who give birth just before and just after the cutoff date of April 1, 1993, provided they cannot manipulate the assignment variable (i.e., the eligibility to a longer maternity leave) near the known cutoff of April 1, 1993, determined by the reform. This is because households cannot choose the birth date of their children exactly, and a difference of just a day can be crucial to be entitled to receive a longer leave or not. If women instead could delay births to benefit from a longer leave, our identification would not be valid. For this purpose, we restrict the sample of analysis to women who gave birth between October 1, 1992 and October 1, 1993, assuming these women were already pregnant at the time of the introduction of the reform and could not have manipulated the timing of their birth to benefit from a longer maternity leave. We check the credibility of this assumption on in Figure 3, which confirms that the monthly density of births for the mothers in the sample is balanced in the 6 months before and after the reform date, which is normalized to 0. Notice also that women who had a child in the six months before April 1, 1993 and had another child in the subsequent six months are very few (less than 1%) and are dropped from our analysis.

An attractive feature of the RD design is that it allows to test the validity of the identification condition. In particular, we check whether the group of women who gave birth 6 months after the reform (treatment group) have similar characteristics to the group of women who gave birth 6 months before (control group). Table 3 shows an equality-of-means test of predetermined characteristics between the two groups. Appendix Table A1 provides a more formal balance test. Even though we find differences that are statistically significant along some of the characteristics (such as age, work experience, tenure and earnings), their magnitudes are not economically important. The more formal test in the Appendix reveals balance across all variables, except mother's age and firm tenure. Using an F-test, we cannot reject the hypothesis that the coefficients of all covariates are jointly equal to zero (p-value=0.129), suggesting the significant differences in age and tenure are not more than what would be expected by chance. Moreover, the point estimates of the differences are small in magnitude and the results hardly change when we include or exclude the characteristics in the RD analysis.

This gives credibility to the identification condition according to which the two groups of women only differ in their eligibility to different maternity leave durations.

Following Hahn et al. (2001), the characterization of our empirical design is given by

$$y_{i\tau}^{(p)} = \alpha + \delta D_i + g(x) + \varepsilon_{i\tau}, \quad \forall x \in [-h, +h] \tag{1}$$

where $y_{i\tau}^{(p)}$ takes value 1 if woman i is a wage earner in percentile p (either in the top decile or top quartile), and 0 otherwise. We estimate (1) using a linear probability model regression, which we repeat at different time points, that is, at different τ years after childbirth over the woman's life cycle, i.e, 2, 5, 10, 15, 20 years after the child's birth. We have an unbalanced panel of mothers, since some may leave the labour market at one point in time (and thus drop out of the registers) and return at a later date.

The assignment follows a deterministic rule: women that fall below the cutoff by giving birth before the reform are assigned to the control group ($D_i = 0$) and can take up to 35 weeks of fully paid maternity leave, while women above the cutoff are assigned to the treatment group ($D_i = 1$) and can take up to 42 weeks of leave. Treatment is defined with respect to the birth date (month and year) of the child, regardless of birth order, but our results are the same when we focus on first births. We estimate the regression equation only for women that gave birth within 6 months around the reform date (between October 1, 1992 and October 1,1993) and refer to this as the bandwidth h. The observation window covers the whole period from the child's birth (in 1992 and 1993) to 2013. We normalize the reform date to 0, and define the running variable, x, as the distance in months between the month of birth of the child and the reform date. In the baseline analysis, the function $g(.)$ is linear in x and is allowed to be different on each side of the cutoff. Because our running variable is discrete, we follow Lee and Card (2008) and assume a parametric functional form around the discontinuity cutoff, clustering the standard errors at the level of distance from the birth cutoff and allowing for robust bias-corrected confidence intervals and inference procedures as in Calonico, Cattaneo and Titiunik (2014) and Calonico, Cattaneo, and Farrell (2018).

The coefficient of interest is δ, which provides the intention-to-treat (ITT) effect of the reform on outcomes for women who gave birth around the reform. In estimating (1), we add further control variables for individual characteristics, including age, schooling, marital status, work experience, and pre-existing number of kids, all measured prior to the reform, in 1992. In Appendix Table A2, we report the estimation results of various specifications around the reform cutoff date, with different choices of the birth window in months. Using the same specification, we have examined all the parental leave extensions preceding the 1993 reform. For space limitation, we do not report the results from that analysis. This is elaborated in Corekcioglu, Francesconi, and Kunze (in preparation (b)). Importantly, the main results discussed below are broadly similar across all reforms.

V. Results

Before presenting the main estimation results, we provide evidence that the distribution of births is continuous around the cutoff date or the 1993 reform. Figure 3 graphically confirms this, showing the density of births within one month bins, 6 months around the reform. It turns out to be the case too for all the preceding reforms from 1988 to 1992. (These results are available upon request.) In the presence of a discrete running variable, we cannot use the standard McCrary test (McCrary 2008) to test for a break in the density distribution of fertility around the reform and use instead the manipulation test proposed by Frandsen (2017). The results from this test statistically reject the possibility of manipulation around the birth month threshold. Moreover, if families time date of birth or change eligibility status in response to the reform, then we would expect to see changes in the distribution of pre-determined observable characteristics of the parents around the reform dates. We also reject this possibility.

Figure 4 shows the unconditional means of the outcomes proportion female in top positions (both $y^{(10)}$ and $y^{(25)}$) by distance (in months) of childbirth from April 1993, within 6 months of the reform. The vertical line in each panel represents the reform cutoff date, whereas the upper and lower lines indicate the 95% confidence interval around a fitted line of the main estimates. We report both short- and long-run impacts, looking at the outcomes 2, 5, 10, 15, and 20 years after the 1993 reform. While the confidence intervals before and after the reform overlap for some of the years at different points of women's lifecycles, the figure suggests an overall negative effect of the reform at the top, which becomes more pronounced as time elapses.

Table 4 summaries our main RD estimates showing the ITT effects, δ in our regression specification (1), which are obtained from regressions within a 6 months birth window around the reform cutoff date. The top panel reports the estimates for women to be in the top decile of the within-firm earnings distribution, and the bottom panel shows the estimates of being in top quartile. We illustrate the impact of the 1993 reform at different points of women's careers focusing on 2, 5, 10, 15, and 20 years after birth (and after the 1993 reform). The years involved therefore are 1995, 1998, 2003, 2008, and 2013, respectively. Models with alternative bandwidths (both smaller and greater than 6 months) lead to relatively similar results, which are therefore not discussed here but are reported in Appendix Table A2.

In 1998, five years after birth and when the women in the sample are approaching 35, mothers exposed to the 1993 reform are 1 percentage point *less* likely to be in the top earnings decile within their organisations. Looking at Figure 2, this is a substantial impact, corresponding to a reduction in the probability of being in top-pay positions by more than one third. The same qualitative result, albeit not statistically significant at conventional levels, emerges when we consider women in the top quartile of the earnings distribution, where a broader set of women in high-ranked positions may be represented if they are systematically paid less than their male colleagues. Although statistically insignificant, the effect is always negative after 2, 10, 15, and

20 years, irrespective of whether we look at the top decile or the top quartile of the within-firm pay distribution.

These results are consistent with the notion that long work interruptions depress wages over women's work cycle. They may also reflect highly nonlinear pay structures at the very top, whereby a lower labour market involvement due to early child care may end up with large wage penalties. With highly paid mothers on average taking their full maternity leave package (see Figure 1 and Table 2), top earnings women may suffer a wage penalty if their human capital depreciates more rapidly or if their pay structure is tightly linked to hours worked. The estimates are also in line with the notion that, as a result of lengthening paid leave, employers may move away from highly paid female workers in their hiring strategies. Identifying the separate impact of each of these explanations is important, but goes beyond the scope of the current paper.

The estimates reported in Table 4 are robust to different specifications of the bandwidth h. They are also robust to the exclusion of all covariates, to a quadratic specification of the $g(.)$ function, and to the estimation of a 'donut' RD, in which we eliminate women giving birth in March or April 1993, one month before and one month after the reform, respectively. Overall, therefore, the parental leave extension does not seem to have delivered a greater representation of women at the top of the earnings distribution within Norwegian firms. If anything, it might have reduced it.

VI. Pointers for Policy and Further Research

Paid maternal leave has gained increased prominence in policy debates worldwide as mothers have increasingly entered the labour market in the past few decades in most advanced economies. As an example, looking at OECD countries, the average number of weeks of paid leave has risen from 17 weeks in 1970 to 51 weeks in 2018 (OECD 2020; see section PF2.5). At the same time, we observe that even very family-friendly countries, such as Norway, struggle to inject more gender diversity in top positions across the whole economy, including central banks, CEOs, supervisory boards, and academia. In this paper, we have asked whether parental leave policy affects women's success within their organisations.

From the existing empirical literature, two different views have dominated the debate. On the one hand, one view recommends generous and long government-mandated paid parental leave, as it is the case in many countries in Europe and, recently, Canada. This view emphasises the contribution of family policies to the goals of gender equity and child development, enabling women to combine careers and motherhood, and altering social norms regarding gender roles.

The adduced evidence in favour of this view is that a long maternity leave, when available, is intensely used by all working mothers. Another consideration is related to child well-being, with evidence suggesting that maternal employment seems to have adverse effects on early cognitive and behavioural child outcomes, especially if employment is resumed full time in the first year of the child's life (Brooks-Gunn, Han, and Waldfogel, 2002; Berger, Hill, and Waldfogel, 2005), but also in pre-school years (Ermisch and Francesconi, 2013). On the positive side, instead, Dustmann and Schönberg (2012) find no effect of long leave on child outcomes, while Ruhm (2000) and Tanaka (2005) provide evidence that infant mortality rates decrease, and child health improves.

On the other hand, an alternative view recommends shorter leaves, as it is the case in the US, the UK, Australia, Japan, Switzerland, and Spain. The arguments underpinning this view warn that family-friendly policies may become an impediment to women's careers because of the loss of work experience and the higher costs to employers that hire women of childbearing age. Some of the evidence used to support these arguments emphasises that shorter leaves, less than 12 weeks, have positive labour market effects, primarily in terms of increased return to work of mothers, and increased probability to return to the same employer (Waldfogel, 1998; Baker and Milligan, 2008). Some more recent study, however, finds that although the introduction of paid parental leave in the US led to increased leave coverage and leave taking, it did not have significant effects on women's wages, earnings, employment and promotions.This has especially been found among women in low-income families, possibly because of the very limited leave length in the United States (Kerr, 2016).

In our analysis, we show for Norway, one country with generous paid parental leave, that women are still underrepresented at the top. We use as a measure of being in the top 10% or 25% of one's own firm pay distribution. The repeated expansions of parental leave, enacted during the period from 1987 to 1993, took place at a time when very few women were among top earners. Representation of women at the top, however, has been gradually increasing over time. An increasing fraction of female top earners have children, albeit still less than men, indicating that in Norway some women may not have to choose between family and career, but can have both. The parental leave policy might have facilitated this process. But we do not find any strong evidence of this.

Women exposed to a longer leave are *not* more likely to reach top positions within their organisations over the whole of their work cycle than women exposed to a shorter leave. If anything, they are less likely to do so. This holds true not just for the latest expansion in 1993 as illustrated in this study, but also successively for each of the earlier reforms introduced every year between 1987 and 1992. Our findings supplement previous findings on Norway. They echo the null results on parental earnings, fertility, and on child school outcomes found by Dahl et al. (2017) on all women. They also chime well with the results by Cools, Fiva, and Kirkebøen (2015), suggesting no increase in mothers' labour market attachment.

One could take our results to suggest that a reduction in parental leave duration or generosity in Norway and in other countries such as Germany, Austria, and Sweden should be beneficial. This is not our conclusion. In fact, even without such a reduction, there might be room for considerable improvements of these systems along a number of other dimensions that require a deeper, more nuanced economic understanding. These dimensions, which are likely to sparkle a new policy debate and future research, can be articulated as follows:

(a) How can government-funded paid parental leave systems be designed without hindering female workers in reaching (and staying in) top-earning positions? Our results are consistent with the notion that long work interruptions can lead to human capital depreciation and this, in turn, is likely to affect wage growth more for highly paid women. We argue this is an important starting point, which policy makers and economic agents (workers and firms alike) should keep in mind.

(b) Although human capital is important, is it the only wage determinant at play? An issue associated with the high depreciation of human capital is the notion of a wage structure that is nonlinear in hours (Goldin and Katz, 2016). If high-earning professions (e.g., lawyers and business consultants) are characterised by highly nonlinear pay structures, then women in such professions will face a motherhood penalty that is substantially larger than mothers in other professions, while fathers do not face any penalties at all. Goldin and Katz (2016) document the successful experience of the pharmacy profession in the United States, where technological changes have led to increased substitutability among pharmacists and growth of pharmacy employment in retail chains and hospitals. With the advance of robotisation and the growing role of artificial intelligence in traditionally white collar jobs, are similar technological changes possible in other sectors and other professions? Would flexible working help working mothers, especially those in high-pay jobs and in the aftermath of a major shock like the COVID-19 pandemic? This is something to be seen in future research.

(c) The family itself may be called to adjust to new challenges. In particular, what is the intrahousehold allocation of chores around childbirth? If parental leave policies aim to give parents an effective way to look after themselves until the first birthday of their babies, then it seems desirable to consider better incentives for fathers to take greater leave in order to achieve a more equal division between parents. Despite the null results found by Ekberg, Eriksson, and Friebel (2013) and Cools, Fiva, and Kirkebøen (2015), according to which (short) paternity leaves do not alter the traditional allocation of parents' labour supply, more substantial extensions to the father's quota could have longer term effects on the division of household work related to childcare. Kotsadam and Finseraas (2011), for example, report a more equal division of household tasks as a result of the introduction of paternal leave in Norway. This point underlines also the importance of social norms and attitudes toward gender roles in the home and in the market, which often accompany – or even anticipate – the introduction and extension of parental leave rights.

(d) Following directly from the previous point, government-funded paid parental leave cannot be seen separately from publicly provided childcare services. Mothers' return to employment cannot be seamless if good-quality free (or affordable) childcare is not available from the start of a child's life, or when parental leave is exhausted. Even in Norway, childcare usage for the youngest has been extremely low until the early 2000s. Besides human capital depreciation, this could be another mechanism behind the negative results we find in our analysis. If families cannot rely on public childcare services, they must turn to private provision, which needs the good functioning of a competitive market place. When children are older and already in school, the organisation of school times is equally essential, especially if children need afternoon care and homework help. The Norwegian system of half-day schools and afternoon school care only until the child is aged 11 does not set a good example and could be in part another mechanism behind our results, especially those that emerge in the longer term. These observations emphasise the importance of coordination among non-parental public and private care providers, primary and secondary schools, as well as employers, which lead us to our next point.

(e) What is the role of firms in shaping family-work balance through parental leave? In countries like Norway, all firms are obliged by law to grant parental leave. Yet, so far, we know very little from large representative employer-employee matched data sets how firms respond to parental leave policies. Thomas (2017) finds evidence that US companies are hesitant to invest resources in the careers of women who they believe are likely to leave to raise families. For the United States, Goldin, Kerr, and Olivetti (2020) find that firm-provided parental leave has greatly increased in the last two decades, but even the top-of-the-line firm in the US provides fewer fully paid parental weeks than does the median OECD nation. Norway is well above the OECD median and mean. But even for Norway, Corekcioglu, Francesconi and Kunze (in preparation (b)) find that firms tend to move away from recruiting young women or young mothers and are less likely to promote them.

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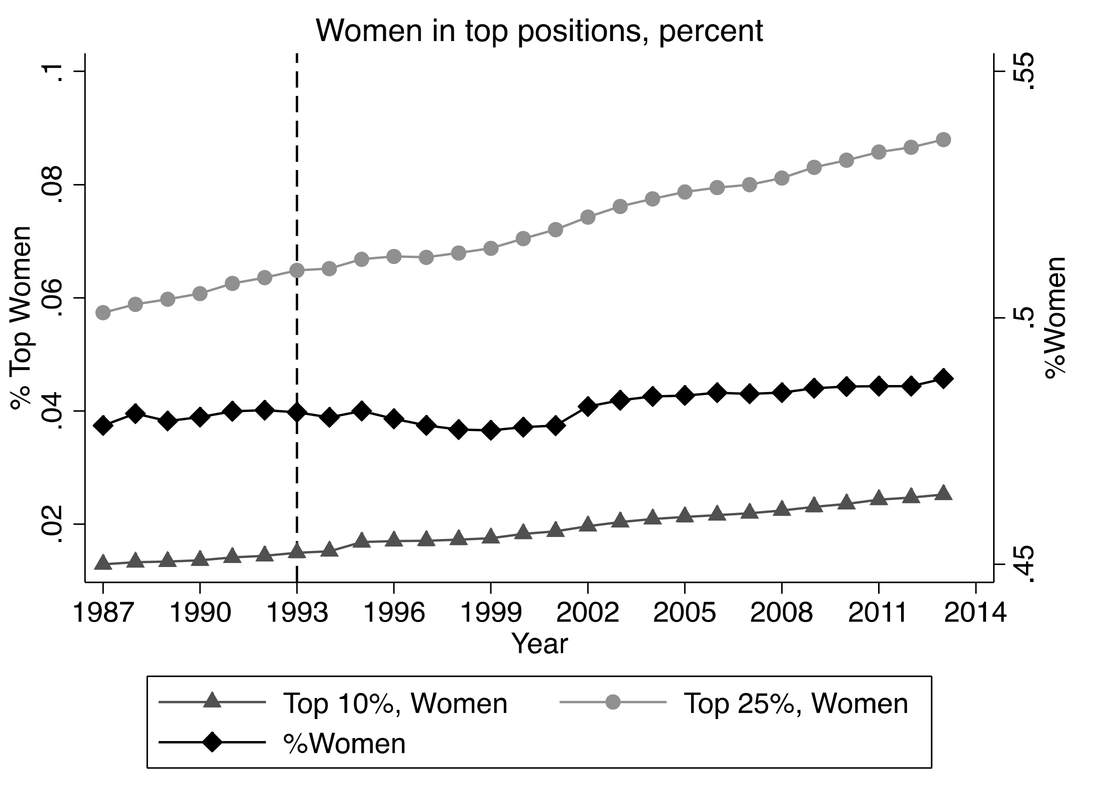
Tables and Figures

Figure 1. Duration of maternity leave around the 1993 reform

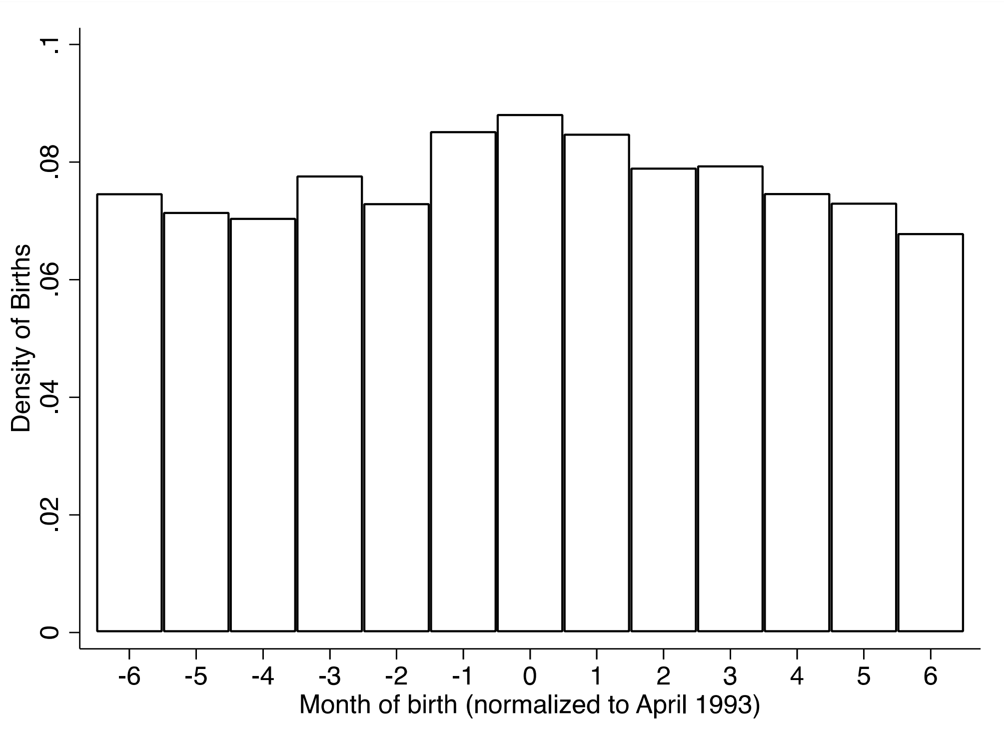
Notes: Each dot shows the average number of days of leave for mothers in one month bins based on the birth month of the child. The dashed vertical line denotes the reform cutoff date, April 1993, which has been normalized to zero. The solid lines are fitted separately on each side of the cutoff, without any controls. Grey lines indicate the 95% confidence intervals around the fitted lines.

Figure 2. Proportions of women in top earnings decile and quartile, and proportion of women in the workforce



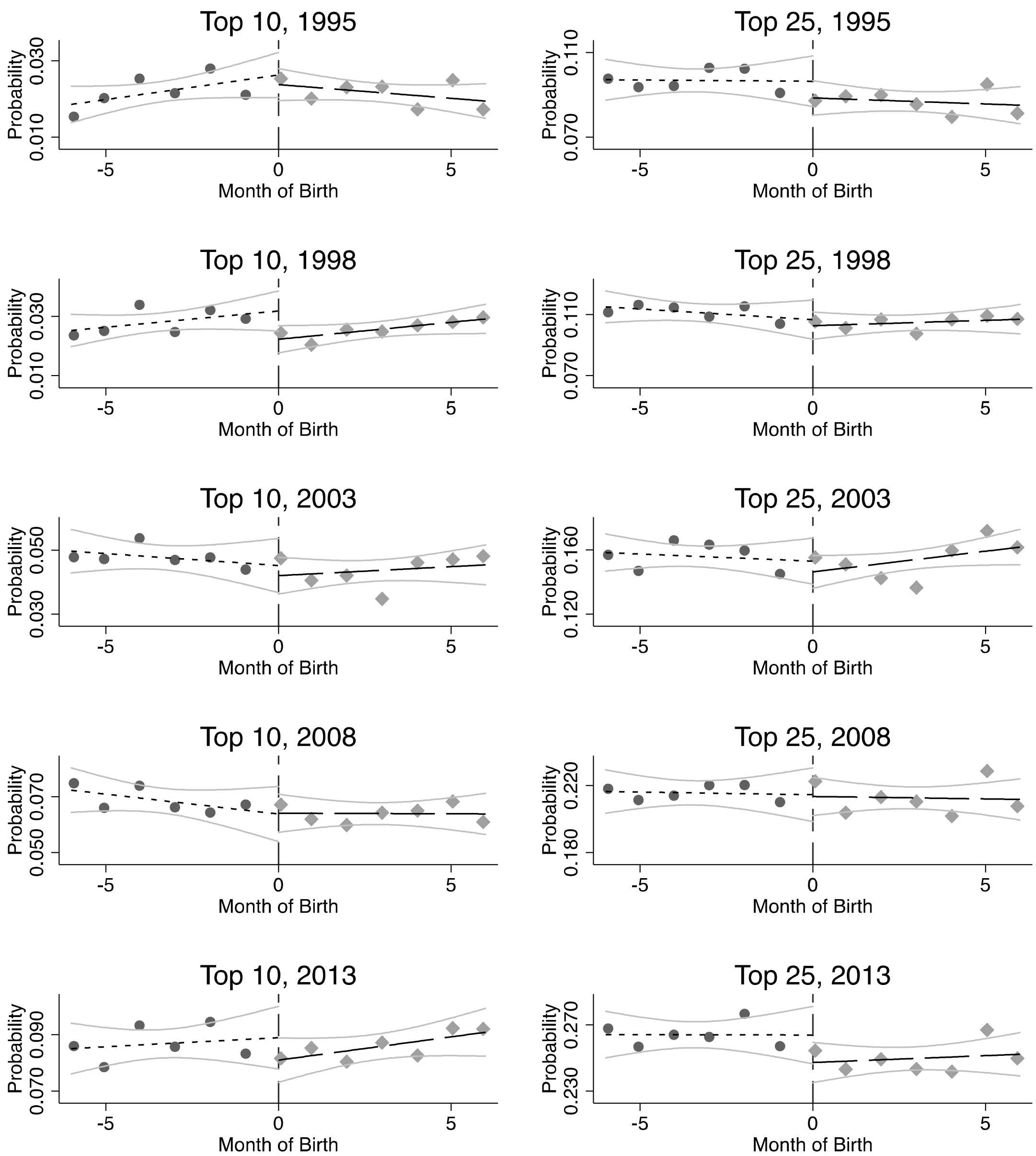
Notes: The left vertical axis refers to the proportion of women in the top decile and quartile of the earnings distribution within firms in any given year. The vertical axis on the right refers to the proportion of women in the workforce.

Figure 3. Distribution of births around the 1993 Parental Leave reform



Notes: Histogram shows the distribution of births by month, 12 months around the reform. The window shown here is restricted to 6 months on each side of the reform. The reform month, April 1993, is normalized to zero.

Figure 4. Women in top earnings decile and quartiles



Notes: Each dot displays the unconditional mean of the relevant outcome variable, within one month bins based on the birth date of the child. The outcomes are the probability of being in the top decile (left panel) and quartile (right panel) of the within firm earnings distribution 2, 5, 10 , 15 and 20 years after childbirth. The dashed lines indicate the reform cutoff date, April 1993, which is normalized to zero. The window is restricted to 6 months on each side of the reform. A regression line without controls is fitted separately on each side of the cutoff, and the lateral lines represent the 95% confidence intervals.

Table 1. Overview of the parental leave reforms in Norway, 1977-1993

Date of the reform	Weeks of leave	Compensation	Maternal Quota	Paternal Quota
01.07.1977	18	100 %	3+6 weeks	
01.05.1987	20	100 %	3+6 weeks	
01.07.1988	22	100 %	3+6 weeks	
01.04.1989	24(30)	100 (80)%	3+6 weeks	
01.05.1990	28(35)	100 (80)%	3+6 weeks	
01.07.1991	32840)	100 (80)%	3+6 weeks	
01.04.1992	35(44.4)	100 (80)%	3+6 weeks	
01.04.1993	42(52)	100 (80)%	3+6 weeks	4 weeks
01.07.2005	43(53)	100 (80)%	3+6 weeks	5 weeks

Source: The Social Insurance Act, see https://lovdata.no/nav/rundskriv/r14-00.

Table 2. Characteristics of top men and top women

	Women, Bottom 90	Women, Top 10	Men, Top 10	Difference (1)-(2)	Difference (2)-(3)
Year=1992					
Age	32.644	34.201	34.645	-1.557***	-0.444***
Education	13.27	14.969	14.566	-1.699***	0.403***
Marital Status	0.569	0.507	0.66	0.062***	-0.153***
Number of Kids	1.41	1.222	1.59	0.188***	-0.367***
Age, youngest child	6.281	7.399	5.27	-1.118***	2.13***
Age at first child	24.345	25.115	26.745	-0.771***	-1.63***
Work experience	5.8	6.206	6.819	-0.406***	-0.613***
Tenure	3.48	3.343	4.301	0.137***	-0.958***
PL take-up	0.82	0.841	0.017	-0.021	0.824***
PL duration, days	268.392	252.189	82.469	16.203***	169.719***
Number of Persons	292,353	7,504	28,372	292,353	35,876
Year=2013					
Age	52.554	52.322	52.49	0.232***	-0.168***
Education	13.248	15.097	14.801	-1.85***	0.297***
Marital Status	0.59	0.611	0.708	-0.021***	-0.096***
Number of Kids	2.042	2.106	2.222	-0.064***	-0.116***
Age, youngest child	21.435	20.242	17.779	1.194***	2.462***
Age at first child	25.706	26.713	28.944	-1.008***	-2.231***
Work experience	23.583	23.524	24.35	0.06	-0.826***
Tenure	9.082	8.805	10.03	0.277***	-1.225***
Number of Persons	361,206	26,733	55,251	361,206	81,984

Notes: Figures are calculated for women and men born between 1952 and 1967 and observed in 1992 (top panel) and 2013 (bottom panel). The figures in the first column refer to all women except those in the top earnings decile within their firms. The figures in the second and third column refer only to women and men in the top earnings decile within their firms, respectively. The fourth column compares the means in the first and second column, while the fifth column compares the means in the second and third column, reporting the difference and its statistical significance.
*** p<0.01, ** p<0.05, * p<0.1.

Table 3. Pre-determined characteristics of mothers who gave birth before and after the reform

	Treatment Group	Control Group	Difference
Year=1992			
Age	29.139	29.503	0.363***
Education	13.619	13.63	0.011
Marital Status	0.496	0.548	0.052***
Number of Kids	0.822	0.806	-0.015*
Work experience	5.377	5.501	0.124***
Tenure	3.22	3.27	0.051*
Log(earnings)	12.131	12.127	-0.004
Top10 Earner	0.025	0.025	0.001
Top25 Earner	0.125	0.125	-0.001
Number of Women	13,507	11,186	24,693

Notes: The figures (measured in 1992) are means computed for all women who gave birth within a year of the April 1993 reform and were at work in 1995. "Treatment Group" consists of women who gave birth between 1 April 1993 and 30 September 1993, while "Control Group" consists of women who gave birth between 1 October 1992 and 31 March 1993, regardless of birth order. The third column ("Difference") compares the means in the first and second column and reports the difference and its statistical significance.
*** p<0.01, ** p<0.05, * p<0.1.

Table 4. Main Results: Probability for women being top earners

Years after birth	2	5	10	15	20
Outcome year	1995	1998	2003	2008	2013
	Outcome: Likelihood, Top 10% Earner				
Reform93 Effect	-0.002	-0.009***	-0.002	-0.000	-0.008
	(0.004)	(0.002)	(0.003)	(0.002)	(0.006)
Mean	0.04	0.04	0.05	0.06	0.07
Birth window(months)	6	6	6	6	6
Observations	25,444	24,887	26,461	27,389	26,763
	Outcome: Likelihood, Top 25% Earner				
Reform93 Effect	-0.006	-0.003	-0.004	-0.001	-0.017
	(0.006)	(0.004)	(0.009)	(0.008)	(0.010)
Mean	0.15	0.16	0.19	0.21	0.23
Birth window(months)	6	6	6	6	6
Observations	25,444	24,887	26,461	27,389	26,763

Notes: The coefficients compare the likelihood of being in the top earnings decile (top panel) or quartile (bottom panel) within the firm for women that gave birth within 6 months prior to the reform date and hence were unaffected by the 1993 parental leave extension, to those that gave birth up to 6 months from the reform date and were exposed to the parental leave extension. Columns (1)-(5) display the intention-to-treat effect 2, 5, 10, 15, and 20 years after the reform. All regression include control variables for age, age squared, schooling, marital status, work experience, and pre-existing number of kids, all measured prior to the reform, in 1992. Standard errors are reported in parentheses. The coefficients are estimated using a linear regression discontinuity model and triangular weights.
*** p<0.01, ** p<0.05, * p<0.1.

Table A1. Covariate balancing tests for characteristics of the mothers

	Age	Education	Married	Parity	Work Exp.	Tenure	Log(earnings)	Top10 Earner	Top25 Earner
Reform93 Effect	-0.269**	-0.006	-0.017	-0.009	-0.141***	0.084*	-0.024*	-0.004	-0.009
	(0.067)	(0.047)	(0.008)	(0.011)	(0.028)	(0.034)	(0.009)	(0.003)	(0.011)
Birth window	6	6	6	6	6	6	6	6	6
Observations	24,703	24,693	24,703	24,703	24,703	24,703	24,703	24,703	24,703
F-test p-value= 0.411									

Notes: All coefficients are estimated using a linear regression discontinuity model and triangular weights, using women giving birth during the birth window 6 months before and 6 months after the reform, and are working in the post-reform year 1995. All covariates are measured in the year prior to the reform, 1992. Standard errors are reported in parentheses. The bottom row reports the *p*-value from a test of joint significance.

*** p<0.01, ** p<0.05, * p<0.1.

Table A2. Robustness to specification and bandwidth choice: Outcome woman in top 10 %

Outcome year	1995	1998	2003	2008	2013
No Controls					
Reform93 Effect	-0.002	-0.009**	-0.002	0.000	-0.007
	(0.004)	(0.002)	(0.003)	(0.003)	(0.005)
Outcome Mean	0.04	0.04	0.05	0.06	0.07
Birth window (months)	6	6	6	6	6
Observations	25,444	24,887	26,461	27,389	26,763
Quadratic					
Reform93 Effect	0.007*	-0.005	0.007*	-0.002	0.002
	(0.003)	(0.004)	(0.003)	(0.004)	(0.005)
Outcome Mean	0.04	0.04	0.05	0.06	0.07
Birth window (months)	6	6	6	6	6
Observations	25,444	24,887	26,461	27,389	26,763
Donut RD					
Reform93 Effect	-0.007*	-0.011**	-0.014*	-0.001	-0.022**
	(0.003)	(0.003)	(0.006)	(0.004)	(0.005)
Outcome Mean	0.04	0.04	0.05	0.06	0.07
Birth window (months)	6	6	6	6	6
Observations	18,879	18,484	19,624	20,332	19,847
3-month window					
Reform93 Effect	0.001	-0.010*	0.004	-0.002	-0.002
	(0.004)	(0.003)	(0.002)	(0.002)	(0.005)
Outcome Mean	0.04	0.04	0.05	0.06	0.07
Birth window (months)	3	3	3	3	3
Observations	14,431	14,110	14,996	15,561	15,226
2-month window					
Reform93 Effect	-0.003	-0.003	0.006*	-0.003**	0.011**
	(0.002)	(0.002)	(0.002)	(0.001)	(0.001)
Outcome Mean	0.04	0.04	0.05	0.06	0.07
Birth window (months)	2	2	2	2	2
Observations	10,198	10,198	10,878	11,236	10,979
9-month window					
Reform93 Effect	0.000	-0.006*	-0.004	-0.002	-0.006
	(0.003)	(0.002)	(0.003)	(0.002)	(0.004)
Outcome Mean	0.04	0.04	0.05	0.06	0.07
Birth window (months)	9	9	9	9	9
Observations	36,749	36,064	38,345	39,775	38,856

Notes: The coefficients are estimated using a linear regression discontinuity model and triangular weights using women giving birth 6 months before and 6 months after the reform. The top three panels report results from alternative specifications: excluding the controls, including a quadratic control function of the running variable, and a donut regression discontinuity model where births in April and May 1993 are excluded. The bottom three panels report results modifying the bandwidth: comparing women giving birth three, two and nine months before and after the reform date, respectively. Columns (1)-(5) display the ITT effect over time 2, 5, 10, 15 and 20 years after the reform. All regression include control variables for age, age squared, schooling, marital status, work experience, and pre-existing number of kids, all measured prior to the reform, in 1992. Standard errors are reported in parentheses.

*** p<0.01, ** p<0.05, * p<0.1.