The Signaling Role of Parental Leave

Linh T. Tô*

31 December 2018

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Abstract

This paper examines the signaling role of workers' parental leave choices theoretically and empirically. Human capital theory predicts that wages should depend on the absolute duration of time out of work. However, using administrative data from Denmark, I find that women who take a given amount of leave earn higher wages after returning to work if other mothers, in the same childbirth cohort, take longer leaves. The importance of relative leave duration suggests the possibility that leave duration choices signal labor market preferences. I develop a model which posits that firms infer private information about commitment to work through their choice of forgoing paid leave to return to work early. The model delivers distinct predictions of the signaling channel of parental leave when there is an exogenous change in the maximum allowed paid leave duration. I test these predictions using unanticipated leave extension policies in Denmark. I show that changing the equilibrium has indirect effects in changing the signals that are sent by all workers, and also has direct effects in terms of labor market consequences. Consistent with the model's predictions, a leave extension causes infra-marginal mothers, whose leave would not have been constrained by the previously lower maximum, to take longer leave. For mothers for whom the previously lower maximum would have been binding, signaling contributes to a divergence in wages due to the information that their choices convey upon a leave extension. The paper provides direct evidence of persistent labor market consequences of signaling in a context in which signaling operates during the course of one's labor market experience.

*Department of Economics, Harvard University. Email: linhto@fas.harvard.edu. I am grateful to my advisors Edward Glaeser, Claudia Goldin, Lawrence Katz, and Elie Tamer for their continued advice and support. I thank Henning Bunzel, Peter Christensen, Jesper Eriksen, Henry Farber, Yana Gallen, Benjamin Golub, Jerry Green, Nathaniel Hendren, Kevin He, Charles Hodgson, Maria Humlum, Myrto Kalouptsidi, Kevin Lang, Robin Lee, Rune Lesner, Ariel Pakes, Theodore Papageorgiou, Mads Schnoor, and Neil Thakral for helpful conversations.

1 Introduction

Despite women's narrowing the gap with men in labor force participation and surpassing them in educational attainment, gender differences in labor market outcomes persist. The persistence suggests an important role for motherhood in explaining the gender wage gap and for labor market policies such as job-protected paid parental leave in addressing gender inequality.[1] The consequences of parental leave policies depend not only on the benefits workers receive but also on the information that firms infer from workers' choices. Due to the uncertainty that firms face about worker effort and persistence, the extent to which a worker makes use of these benefits conveys information about the value of employing that worker. The use of such information by firms then contributes to statistical discrimination. Programs meant to provide workers with benefits to mitigate inequality between men and women can thus give rise to inequality among women.

This paper analyzes the signaling role of workers' parental leave decisions in determining earnings and wages. The model posits that a worker's decision to forgo paid leave serves as a costly signal to the firm of the future value of employing that worker. The empirical finding of a negative relationship between parental leave duration and subsequent labor market outcomes supports this interpretation. In an argument that parallels that of the classic signaling framework by Spence (1973), when private benefits from leave taking are correlated with unobserved productivity or commitment, workers who face lower costs of forgoing leave earn higher wages in equilibrium. The signaling channel of paid leave makes distinct predictions about the relationship between wages and firms' beliefs about workers. Testing these predictions relies on an exogenous shock that leads to changes in equilibrium choices.

That exogenous shock is found in a series of parental leave policy changes in Denmark. Leveraging administrative longitudinal data from Denmark together with

[1] Altonji and Blank (1999) summarize the earlier literature that explains the differences in labor market outcomes by race and gender focusing on pre-market human capital differences and discrimination. Recent studies identify that the biggest divergence between men and women in advanced countries occurs after childbirth, including in Sweden (Angelov et al., 2015), Denmark (Kleven et al., 2018), Germany (Adda et al., 2017), the United States (Bertrand et al., 2010) and the United Kingdom (Kuziemko et al., 2018), among others.

these policy changes provides an ideal empirical setting for testing the predictions of the signaling model. Denmark has a long history of federally provided paid parental leave, with several unanticipated changes in the maximum benefit duration. The data track workers throughout their entire careers and contain information on individual leave taking duration, a variable often overlooked in studies evaluating policy changes. The long-standing high-quality public provision of child care in Denmark means that mothers' tradeoffs stem primarily from labor supply preferences rather than constraints, allowing for an analysis of the channels that the model highlights.

Section 2 of the paper presents an event-study framework that elucidates the correlation between leave choices and labor market outcomes. The results show that after childbirth, women who take longer leave relative to others within the same year experience greater reductions in labor supply and hourly wages conditional on working, despite having similar pre-birth career trajectories. The "child penalty" (Kleven et al., 2018) thus exhibits substantial heterogeneity across mothers based on *relative* leave duration.[2] Whether this heterogeneity arises due to differential labor-supply preferences or a causal effect of leave taking, firms can infer information from the leave duration decisions of mothers.

Section 3 formalizes the link between wages and firms' beliefs about the future value of employing a worker. Any correlation between workers' private benefit of leave taking and productivity or commitment leads wages to depend on firms' beliefs, which in turn depend on observed leave choices in equilibrium.[3] This induces workers to take less leave than they would if firms' beliefs (and hence wages) did not respond to observed leave taking decisions. The model highlights that viewing labor market policies solely in terms of the direct benefits to workers ignores the interplay between firms' and workers' choices and its implications.

[2] Albrecht et al. (1999) highlight that *absolute* maternity leave duration looking at cohorts of mothers over time may not be associated with lower wages. The difference between absolute versus relative leave duration is consistent with the signaling channel.

[3] A correlation between wages and leave duration might arise even without directly assuming that less productive workers prefer take longer leave. For example, if firms face a cost to hiring a new long-term worker, then more committed workers generate higher rents for the firm. Thus, wages depend on firms' beliefs insofar as leave duration provides information about a worker's disutility of effort.

To bring the theory to the data, I derive testable predictions of the model corresponding to an increase in the maximum allowed leave duration. A policy change that alters equilibrium choices can then help distinguish the signaling channel. Absent such a change, a negative relationship between leave taking and wages would not necessitate asymmetric information. For example, a worker may simply choose not to take longer leave due to potential losses in human capital and therefore future wages. In response to a shock that expands the choice set to include longer leave options, the symmetric information model predicts that workers who would choose strictly less than the maximum leave duration would not change their behavior.[4] In contrast, if a worker's decision to take less than maximum leave conveys private information to the firm, then the model predicts a *ripple effect*: When less valued workers take longer leave in response to the change, higher valued workers would be enabled to take longer leave as well. The asymmetric information model thus predicts that workers for whom the maximum allowed duration does not impose a binding constraint would alter their choices.

I further establish the distributional consequences of paid leave on wages and earnings due to signaling. Understanding these effects requires an analysis of *partial pooling equilibria*, which arise due to the mandated maximum leave duration.[5] In particular, I focus on the class of equilibria consistent with the data for mothers in Denmark, in which less valued workers pool by choosing the maximum leave allowed. I show that an increase in the maximum leave duration reduces the fraction of mothers pooling at the maximum. This prediction leads to shifts in the distribution of wages for different groups. Workers who would have pooled at the maximum leave duration before, but not after, a leave extension would be able to reveal their type as a result of the change. Because these workers are better types than the average worker who would have pooled before the leave extension, the model predicts that their wages increase. The workers who pool at the maximum after the extension are those who generate less value to the firm, and the model predicts their wages decrease. An

[4]The implicit assumption is that preferences for longer leave can have at most one local optimum.

[5]Under a fully separating equilibrium, even if the distribution of leave choices shifts, wages and earnings remain unchanged because a such equilibria reveal information perfectly.

increase in the maximum paid leave duration can thus increase earnings inequality among women.

Section 4 carries out the empirical tests using two parental leave reforms in Denmark: one in 1984 and another in 2002. Mothers could not have anticipated either policy change until within three months of the effective dates, and the implementation used sharp birthdate cutoffs. For both reforms, mothers and fathers could share the extended leave amount, but in practice the incidence of shareable leave falls almost entirely on mothers. The signaling channel may also help understand fathers' choices not to take up shared leave, a point of discussion that I revisit in the conclusion.

Section 5 discusses how the predictions are revealed in the data. Both policy changes result in a shift in the distribution of leave taking. Furthermore, for mothers with childbirths around the reform window who take a given below-maximum level of leave, the post-reform mothers earn higher wages. To test the distributional wage effects of parental leave, I define subgroups of mothers based on their pooling status. Based on the child's date of birth, I observe whether each mother pools at the maximum allowed leave duration. I then match pre-reform and post-reform mothers by their pre-birth characteristics to form the subgroups pre-pooler post-nonpooler (PN) and pre-pooler post-pooler (PP). The data confirm the prediction of the signaling model that these groups experience opposite effects on wages and earnings in response to the leave extension.

Channels other than signaling cannot account for the patterns in the data. The shift in the distribution of leave taking among mothers for whom the leave extension does not relax a binding constraint necessitates an information channel. An alternative explanation based on spillovers would imply that mothers change their behavior after observing that others take longer leave, even if their actions do not convey any relevant information to firms. However, the sharp and persistent change in the leave distribution provides evidence against this view.

In addition, existing work focuses on the role of continued employment and human capital losses due to time away from work, which cannot explain the observed heterogeneous effects by pooling status. Increased employment can result in negative selection of who works and thus a negative estimated effect on average wages. However, the data show that the positive employment effect for the PN group is larger than that for the PP group. Negative selection thus explains neither the absolute increases in wages for the PN group nor the wage gains for the PN group relative to the PP group. Furthermore, the data do not show evidence of a meaningful impact of productivity losses in response to small scale increases in leave taking. While mothers who do not pool at the maximum increase their average leave duration in response to a leave extension (the ripple effect noted earlier), both reforms result in null effects on wages. I further confirm that observed differences in characteristics across the groups do not explain their differences in outcomes.

The most closely related papers examine the role of signaling in labor market investments, namely education, but without any direct empirical evidence on labor market outcomes. Lang and Kropp (1986) and Bedard (2001) test the educational sorting hypothesis against the human capital hypothesis by looking at how actions (i.e., obtaining education) change when the cost of signaling varies. In particular, these papers test whether a change that directly affects one group leads to an equilibrium change in the choices of another group, much like the shift in the distribution of leave taking among mothers for whom the maximum allowed leave duration does not impose a binding constraint in this paper. In contrast to these papers, the current paper further examines effects on wages and earnings and highlights the distributional consequences of signaling.[6] Both Lang and Kropp (1986) and Bedard (2001) exploit cross-sectional differences in the cost of obtaining education, due to differences in the compulsory attendance laws in the former and university access as proxied by the existence of local universities in the latter, whereas the current paper makes use of an exogenous policy change. In addition, while the literature on signaling in education relies on differences in the cost of signaling (obtaining education), the current paper considers differences in the private reward of the signal (benefit to staying at home

[6]Others papers testing the signaling value of educational attainment look at the “sheep-skin” effect of obtaining the degree itself, such as Murnane et al. (2000) and Clark and Martorell (2014). Fang (2006) empirically quantifies the contribution of ability signaling to the college wage premium while noting the reliance on functional form assumptions. The current paper in contrast exhibits directly the consequence of the signaling channel.

at a fixed wage).[7]

Few applications of signaling theory focus on pooling equilibria.[8] The model in this paper relies on the fact that there is an upper bound on the action space, because of the maximum allowed leave duration, which necessitates considering partial-pooling equilibria. Because of private information, expanding the choice set to include longer leave options creates separation between otherwise indistinguishable groups of workers, which directly affects their labor market outcomes. This contrasts with the case of full separation, where signaling functions solely as an unproductive reduction in leave taking with no equilibrium consequences for wages.

The paper also relates to the larger literature on asymmetric information in the labor market. Several of the papers in this literature discuss the role of statistical discrimination in explaining the gender wage gap, summarized by Altonji and Blank (1999) as well as in more recent work by Albanesi and Olivetti (2009) and Gayle and Golan (2012). Landers et al. (1996) describes the "rat race" across workers, in which workers respond to asymmetric information about the propensity to work hard by working inefficiently long hours. Gibbons and Katz (1991) and Kroft et al. (2013) analyze the informational content of types and duration of unemployment, though these events do not involve a signaling action on the part of the workers.

Existing studies on parental leave focus on evaluating the direct average benefits and costs of extending paid leave. The results in this paper agree with the large body of work finding minimal average long-term effects on earnings and wages of these policy changes (see Olivetti and Petrongolo 2017 for a summary). The current paper focuses, however, on a different channel that the existing literature does not discuss, namely that signaling concerns inhibit workers from taking full advantage of paid leave benefits. In the process, mothers with smaller private benefits from leave taking can

[7] Lang and Manove (2011) also use the benefit angle rather than cost. They look at racial differences in educational attainment conditional on ability measured by the Armed Forces Qualification, and point out that blacks invest more in education because the productivity value of the signal is higher for them.

[8] Exceptions include Bernheim and Severinov (2003) and Andreoni and Bernheim (2009) when pooling equilibria represent social norms as well as Kartik (2009) on strategic communication with costly misreporting. The latter notes the benefit of additional insight from comparative statics of partial separation at the cost of the more complex analysis required.

differentiate themselves by taking less leave, contributing to inequality in outcomes among women. Blau and Kahn (2013) and Thomas (2018) offer a complementary perspective, discussing the role of asymmetric information in contributing to statistical discrimination by gender.[9]

Finally, a growing body of work documents the impact of children on women's labor market outcomes. Most notably, using survey data from the United States and the United Kingdom, Kuziemko et al. (2018) suggest that women face uncertainty about the employment costs of motherhood, as their stated attitudes on work-family balance change upon childbirth. The uncertainty mothers face supports the asymmetric information assumption in this paper. If mothers themselves face uncertainty that does not resolve until after childbirth, then forgoing parental leave serves as perhaps the first costly action to convey their private information to firms. This paper also builds on the event-study analyses in Bertrand et al. (2010); Angelov et al. (2015); Chung et al. (2017); Kleven et al. (2018); Hotz et al. (2018); Kuziemko et al. (2018) by adding a new dimension of heterogeneity of the "child penalty" by leave duration and highlighting that a one-time event predicts a lasting impact on subsequent labor market outcomes.[10]

2 Data and patterns of parental leave

The paper uses administrative data from Denmark to examine patterns of parental leave taking from a country with a long tradition of family leave protection. While the framework in this paper is not specific to a particular setting, the empirical analyses

[9]They discuss how paid leave induces women at the margin of employment to select into work, and firms are less likely to invest and promote women in response. The asymmetric information channel in the current paper, by contrast, relies on differential take-up of benefits which serves as a signal.

[10]The finding is consistent with Preston (1997)'s study of men and women in science fields at a public university. Using survey data for 1,700 individuals, she documents that women who report to take on zero percent of child care responsibilities have an earnings premium over similar men, and parents who take on 100 percent of child care earn 29 percent lower salaries than those having no child care responsibilities. Child care responsibilities naturally relate to subsequent labor market outcomes since they may continually and directly require time away from working.

have the advantage of examining the theory using rich historical data covering a long period with multiple policy changes in Denmark. Appendix A provides a brief summary of the history of parental leave programs throughout the period of my data.

2.1 Data

I combine register data administered by Statistics Denmark covering the entire Danish population between 1980 and 2012.

The Social Statistics register (SHSS) provides information on days taken for parental leave between 1984 and 2007, and the Ministry of Employment progress database DREAM further supplements weekly information on parental and child care leave between 1992 and 2012. Parents and children are linked through the birth register, which includes precise birth date information and comprehensive linkage between different generations in the same family.

To evaluate mothers' labor market outcomes, I rely on the extensive information from the Danish integrated database for labor market research (IDA). The data contain comprehensive information about the primary employment in November each year, including detailed worker characteristics such as gender, age, education, experience, tenure, hourly wages, and annual earnings, together with firm and industry information. The IDA data cover the period between 1980 and 2012.

A unique advantage of the Danish register data is its measure of hours worked which is available for the full population. The variable is available starting in 1964 with the introduction of a mandated pension scheme (Arbejdsmarkedets Tillægspension), for which employers are required to contribute for their employees based on weekly hours worked.[11] Due to the early availability of this variable, I also obtain a measure of work experience for the majority of grandparents of children born after 1984.

2.2 Patterns of leave taking

This section provides an overview of the differences among women along the leave-duration dimension, both before and after becoming a mother. Instead of comparing

[11] Kleven et al. (2018) describe the measure in more detail.

between men and women, the analysis focuses on comparing women who choose to take longer parental leave to those who take shorter leave. The analysis explores the similarities in mothers' pre-birth characteristics, while noting that their labor market outcomes are correlated with their leave-taking decision, with longer leave being associated with a larger impact of childbirth.

The sample consists of mothers whose first children are born between 1984 and 2003 and who take at least one day of parental leave.[12] In total, there are 440,605 mothers in the sample, who are alive and reside in Denmark (and therefore appear in the data) for at least four years before and ten years after the event of childbirth. Mothers whose children are born between 1992 and 2001 are under the child care leave scheme in addition to parental leave. In that case, the leave duration measure includes both parental leave and any child care leave that is used before the child turns two.[13]

Leave taking and pre-birth characteristics

I first look at personal and labor market pre-birth characteristics of mothers who take above versus below the median leave length for their first births. Appendix Figure 1 shows the distribution of age, education, log wage, and industry the year before a woman becomes a mother for the first time.

Over this period, the median age of first-time mothers is 27. Their pre-birth wage rate is approximately $20 per hour measured in 2000 USD. Vocational school is the most representative education level among mothers in the sample at 40 percent, with less than 30 percent of mothers having some college or higher degree. The industries that new mothers cover are those with traditionally higher female representation such as education and public services, as well as sales and services.

All four characteristics show similar patterns by leave duration. Mothers whose take shorter leave length are slightly older, more educated, with somewhat higher pre-birth wages, and they are more likely to be in the finance and real estate industry.

[12]The parental leave restriction results in around 90 percent of mothers to be included.

[13]Beuchert et al. (2016) note that many women use child care leave to extend their parental leave, which I confirm in my data.

While these patterns are intuitive, knowing a mother's pre-birth characteristics is not highly predictive of her subsequent leave duration.

Leave taking and labor market outcomes

I turn to examining the post-birth labor market outcomes of mothers by leave duration. For each childbirth cohort between 1984 and 2003, I group first-time mothers into quartiles based on post-birth leave taking time relative to mothers giving birth in the same year. Since the generosity of the leave regime increases over time, I use cohort-level quartiles to examine relative leave taking within a given policy regime. I then compare the outcomes of mothers relative to fathers by leave quartile around the event of childbirth, from five years before to ten years after.

The analysis has the same structure as those considered by Kleven et al. (2018) and Angelov et al. (2015) to study the impacts of children on mothers. I extend the analysis to study the heterogeneity in labor market outcomes across mothers by their choice of leave duration. I use the following regression for mothers and fathers together:

$$Y_{iyt} = \sum_{j \neq -1} \alpha_q^j (\text{event}_{tj} \times \text{female}_i) + \sum_j \gamma_q^j \text{event}_{tj} + \delta_{ga} + \nu_{gy} + X_i \beta + \epsilon_{iyt} \tag{1}$$

Here, the outcome Y_{iyt} for individual i in calendar year y (t years relative to childbirth) is compared across gender g and parental leave duration quartiles q. The dummy variable event_{tj} is equal to 1 when the years-from-birth t is equal to j. The specification takes out non-parametric time trends and life-cycle trends by flexibly controlling for age a and year fixed effects y interacted with gender. The pre-birth controls X_i include education, family status (single, cohabit, or married), and own and spouse income six years before the birth of the first child. The coefficients α_q^t measure the relative difference between women and men, normalized to be 0 in the year just before the first child.

Following Kleven et al. (2018), I convert the estimated level effects into percentage of the counterfactual outcome in the absence of childbirth. This measure removes the unit of the variables and standardizes them. In particular, I adjust $\hat{\alpha}_q^t$ to $p_q^t \equiv \hat{\alpha}_q^t / \mathbb{E}\left[\bar{Y}_{ist} \,\middle|\, t\right]$, where $\bar{Y}_{ist}$ is the predicted outcome when setting the years-from-birth variable t to always be -1.

I consider four labor-market outcomes, total earnings which are zeros for women who do not participate in the labor market, labor market participation rates of all women, as well as hourly wage and total annual work hours conditional on staying in the labor market. Figure 1 plots the adjusted coefficients p_{tq} by quartile group $q \in \{1, 2, 3, 4\}$ over years-from-birth $t \in \{-5, \ldots, 10\}$. The 95% confidence bands are computed using robust standard errors.

Two clear patterns emerge. Prior to becoming mothers, the relative earnings of women compared to men are close to being at the same level of the year before birth throughout the five-year window, independent of their future leave choice. All four groups of women suffer a sharp drop in annual earnings after childbirth, but the trajectories are distinct across the leave-duration quartiles. In particular, women who take longer leave have lower post-birth earnings despite being on a similar earnings path pre-birth compared to women who take shorter leave. Three years after the first child arrives, a woman whose leave duration falls in the fourth quartile has an average drop in earnings of 40 percent, twice as large as that for a woman whose leave duration falls in the first quartile. This ratio after ten years is around 1.5.

Annual earnings can be decomposed into labor market participation, hours worked, and wage rates. Across all three of these margins, leave duration is predictive of labor market outcomes of mothers, with a higher leave length being associated with a career path in which women are either less attached to the labor market or are in positions with lower wages.

The above described heterogeneity across the four leave-duration quartiles is not meant to capture the causal impact of leave choice.[14] Mothers who choose to take longer leave may very well have different labor supply preferences from mothers who take shorter leave. What the analysis confirms however, is that women who look very

[14]The event-study design itself attempts to capture the causal impact of childbirth on labor market outcomes. The identification of short-term outcomes relies on a smoothness assumption. For longer-term effects, Kleven et al. (2018) compare their estimates to two alternative specifications, using placebo mothers, and using a sibling sex mix instrument.

similar prior to childbirth both in terms of personal characteristics and labor market trajectories can have very different career paths along the leave length dimension. The descriptive evidence is thus consistent with a world where parental leave duration reflects post-birth labor market choices, but unanticipated by the firms.

3 A signaling model of parental leave

Informed by the descriptive facts from Section 2.2, this section presents a signaling model of parental leave, where labor market attachment after childbirth is private information which can be revealed through the choice of leave, and firms obtain rents when workers are more committed to the same job. In the spirit of the classic work by Spence (1973) on job market signaling, I consider a pure signaling model and shut down other channels such as the human capital effect of leave taking. In the model, the equilibrium wage function depends on observed productivity and unobserved labor market attachment. The goal is to generate testable predictions that can be brought to the data. The main comparative statics from the model are with respect to changing the action space by relaxing the maximum leave duration. The model therefore departs from the standard Spence (1973) framework with discrete types and instead features continuous types and continuous actions.

3.1 Model

The model microfounds the dependency of equilibrium wages on labor market attachment, which is defined as a parameter that determines the disutility of work for mothers. Similar to the Spence (1973) framework, the model does not allow firms and workers to engage in a contingency contract over an extended period of time. The distinctive feature that causes firms to potentially pay workers a surplus over their marginal productivity is that firms gain rents by saving on hiring costs when an incumbent worker is expected to stay at the firm. In this private information setting, the hiring costs give rise to statistical discrimination based on labor market attachment (Barron et al., 1993; Gayle and Golan, 2012).[15]

Setup

The game involves two sides, firms and workers.

There are infinitely many identical firms in a competitive market. Each firm hires one worker in each period and wage is only binding each period. A firm has to pay a fixed hiring cost τ in the first period when a new worker joins the firm.

Women are identical prior to period 1 but their disutility of effort parameter (type) θ differs starting period 1 when children are born. This is their private information and the firms only know the distribution F_θ with support $[\theta_{\min}, \theta_{\max}]$. Parameter θ scales the effort cost function $C(\cdot)$ which is increasing and convex. Each woman has a fixed observed productivity y.

For simplicity of the notation, I will assume a discount factor of 1 for both the firms and the workers.

Timeline

The model includes 2 periods, starting at period 1 when a worker has her first child, with private information about her disutility of effort. Period 2 serves the role of a terminal period, which effectively reduces the model to a static one, as in Spence (1973).

Period 1 All women work but they can choose to take ℓ parental leave up to $\ell_{\max} \leq 1$. The firm determines the wage schedule $w_1(y, \ell)$ based on observed productivity y and action ℓ and pays at the end of the period. Because we are only concerned with the case of paid parental leave, we assume that there is no cost to the firm when the workers take leave. In particular, the firm gets reimbursed for the leave length that the worker takes, and can simply use the wage portion not

[15] There are other ways to microfound how wages may depend on parental leave as a signal of private information. An alternative model where productivity in future periods is correlated with the private value for parental leave will introduce a direct link between parental leave choice and firms' inferences about future productivity and consequently wages.

paid to the worker to hire a substitute worker at no cost, maintaining the same total productivity.[16] Worker i's utility in period 1 is linearly separable in wages and disutility of effort and defined as $u_{1i}(\ell) = w_1(y_i, \ell) - \theta_i C(1-\ell)$.

Period 2 The firm determines the wage schedule $w_2(y, \ell)$ and pays at the end of the period. Worker i receives a random shock $\epsilon_i \sim F_\epsilon$ which is a uniform distribution, and then chooses to work at the scheduled wage or to receive the outside option of 0. In particular, F_ϵ is $\mathcal{U}[-(y - \tau - \theta_{\min} C(1)), -(y - \tau - \theta_{\max} C(1))]$, chosen as the smallest interval which makes the probability of working or not working nonzero. If the worker works, her utility in period 2 is then $u_{2i}(\ell) = w_2(y_i, \ell) - \theta_i C(1)$.

Equilibrium wage and utility function

The problem over the two periods can be solved using backward induction to derive the equilibrium wage functions and the leave choice condition. Note that there is no uncertainty in period 2 about the actions of the workers after this period.

In period 2, a firm who hires a new worker (y, ℓ) earns profit $y - w_2(y, \ell) - \tau$. Competitive identical firms earn 0 profit and therefore $w_2(y, \ell) = y - \tau$. The incumbent firm pays the market wage but not the hiring cost, so its profit is τ. Taking the firm's action into account, worker i chooses to work if $\epsilon_i > \theta_i C(1) - (y_i - \tau)$ with ex-ante probability $1 - F_\epsilon(\theta_i C(1) - (y_i - \tau))$.

In period 1, a firm who hires a new worker (y, ℓ) has expected total profit $(y - w_1(y, \ell) - \tau) + \tau\Big(1 - F_\epsilon(\hat{\theta}(\ell) C(1) - (y - \tau))\Big)$ given its belief $\hat{\theta}(\ell)$ of the worker's type which determines the competitive wage schedule. The competitive wage schedule is $w_1(y, \ell) = y + \tau\Big(1 - F_\epsilon(\hat{\theta}(\ell) C(1) - (y - \tau))\Big)$. The expected utility of a worker with (y, θ) this period when taking leave ℓ is then $U(\theta, \hat{\theta}, \ell) = (y - \tau) + \tau\Big(1 - F_\epsilon(\hat{\theta}(\ell) C(1) - (y - \tau))\Big) - \theta_i C(1 - \ell)$.

In this simple setting, because the actions of the workers in period 1 do not affect their payoffs beyond this period, $U(\theta, \hat{\theta}, \ell)$ is the *reduced-form utility function*

[16] Brenøe et al. (2018) and Gallen (2018) study the costs to firms and coworkers when mothers take parental leave, and find negligible impacts on employment and wages of coworkers, total wage bill, firm outputs, profits, or closures.

that informs the equilibrium leave choice. The next sections analyze the signaling equilibrium of parental leave taking into account the reduced-form utility function directly.

3.2 Signaling equilibrium characterization

As with other signaling problems, there are multiple Perfect Bayesian Nash equilibria.[17] This section characterizes two classes of signaling equilibria; fully separating equilibria when every type is fully revealed, and partial pooling equilibria when only the least committed types pool at the maximum allowed leave.

Definition 1. A Perfect Bayesian Nash equilibrium (PBNE) of the reduced-form signaling problem consists of two elements, workers' leave strategies given their types $\ell(\theta)$, and firms' beliefs $\hat{\theta}(\ell)$ at every possible strategy $\ell \in [0, \ell_{\max}]$. In equilibrium, workers maximize $U\Big(\theta, \hat{\theta}(\ell), \ell\Big)$ given the firms' beliefs, and the beliefs follow Bayes' rule whenever possible.

A fully separating equilibrium is one in which every type takes a distinct action and employers' beliefs coincide with the true types.

Proposition 1 (Fully separating)**.**

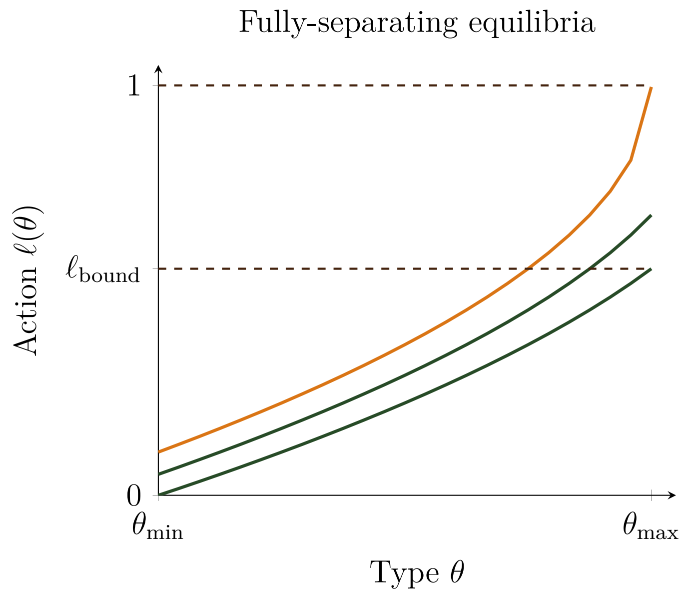
1. *(Characterization) The worker's strategy $\ell(\theta)$ in any fully separating equilibrium solves the differential equation (DE) $\frac{d\ell}{d\theta} = -\frac{U_2}{U_3}$, and the firms' belief is $\hat{\theta}(\ell(\theta)) = \theta$. All mothers take strictly less leave than first-best, except for possibly the least committed type.*

2. *(Existence) A fully separating equilibrium exists if and only if, at the boundary value where the solution to the DE satisfies $\ell^*(\theta_{\min}) = 0$, $\ell_{\text{bound}} \equiv \ell^*(\theta_{\max}) \leq \ell_{\max}$.*

The proof can be found in Appendix B.1. The proof applies the results in Mailath (1987) and Mailath and von Thadden (2013), checking the properties of reduced-form

[17] In an overview, Hörner (2008) associates signaling problems with the literature on equilibrium selection, and screening problems with addressing possible nonexistence of the solutions.

utility function U which allows equating incentive compatibility to solutions of the DE. Among the conditions, monotonicity in type (less attached types benefit more from taking more leave) and single-crossing (a more committed type is always willing to take shorter leave slightly more to differentiate herself from a less committed type) highlight the tradeoffs for mothers.



The diagram illustrates that when fully separating equilibria exist, the most efficient one is the one in which the least committed type takes the full amount of leave. This is also the unique equilibrium in this class that is undefeated in the sense of Mailath et al. (1993).

If the action space is sufficiently small that the DE cannot be satisfied for all types, then there is not enough room for all of the types to separate and hence no fully separating equilibrium. The proof formalizes this constraint and confirms the intuition that when the hiring cost τ is higher, the action space has to be larger to facilitate a fully separating equilibrium.[18]

[18] While this is not a comparative static that I take to the data directly given that hiring costs are unobserved, and there are alternative ways to microfound firms' beliefs, Appendix Figure 1 shows that women in Finance and Real Estate tend to take shorter leave, consistent with the intuition of the result.

I next consider a particular class of partial pooling equilibria that is consistent with the data: a significant fraction of women take the maximum allowed leave.[19] To facilitate doing comparative statics and generating testable predictions, I use the undefeated equilibrium refinement by Mailath et al. (1993) to select a Pareto optimal equilibrium in this class.

Proposition 2 (Partial pooling)**.** *Consider the class of partial pooling equilibria in which the least committed types pool at $\ell_{\max}$ and the other types separate:*

1. *(Characterization) An equilibrium in this class is defined by a cutoff type θ^*, such that $\ell(\theta)$ for $\theta \leq \theta^*$ is a solution of the DE, and $\ell(\theta) = \ell_{\max}$ for $\theta > \theta^*$. For each θ^*, there exist beliefs that sustain the partial pooling equilibrium above as a PBNE.*

   For uniform types, the type θ^* that is indifferent between pooling and separating chooses a leave level on the indifference curve $\phi(\theta)$ which satisfies $U(\theta, \theta, \phi(\theta)) = U\Big(\theta, \frac{\theta+\theta_{\max}}{2}, \ell_{\max}\Big)$.

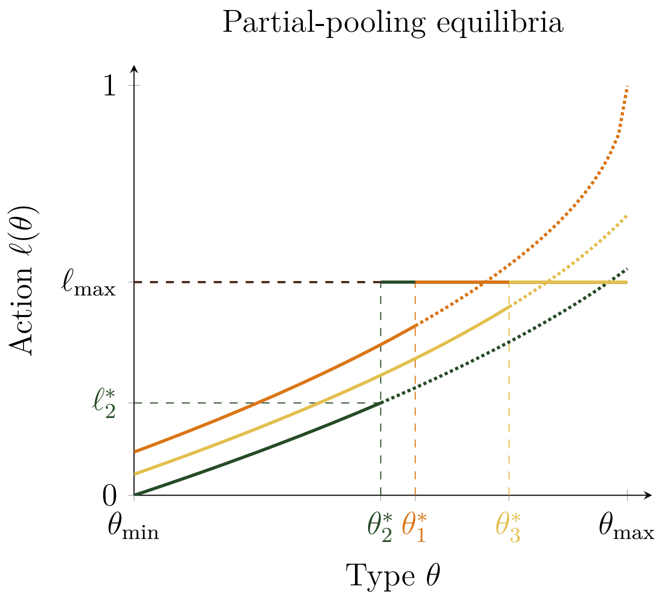
2. *(Refinement) The equilibrium with the most committed type taking the longest possible leave is the lexicographically maximum equilibrium (LME) and is undefeated in the sense of Mailath et al. (1993) under weak assumptions.*

3. *(Property of the refinement) Within this class, the LME strictly Pareto dominates all undefeated equilibria.*

The proof can be found in Appendix B.2. The proof first considers candidates for the LME and shows that the considered equilibrium is the unique LME as long as the most committed type prefers it to the fully pooling equilibrium when all types take maximum leave. Applying Mailath et al. (1993), the LME is undefeated.[20]

[19] This is especially true in the early 1980s, the setting of the tests in Section 4, when there are only 14 to 20 weeks of allowed paid leave.

[20] The belief-based undefeated equilibrium refinement admits more equilibria than the D1 refinement. D1 refinement restricts the equilibrium to one in which only the most committed types can pool (Ramey, 1996), which is outside of the equilibrium class considered. While D1 can be useful to eliminate pooling equilibria when fully separating equilibria exist, it can often be too strong, as acknowledged by Cho and Kreps (1987) in the context of the Spence (1973) model.



The diagram above illustrates three partial pooling equilibria in the considered class, equilibrium 1 (orange), equilibrium 2 (yellow), and equilibrium 3 (green). The types θ_i^* ($i \in \{1, 2, 3\}$) is indifferent between pooling and separating in each of the three equilibria and they all lie on the indifference condition curve $\phi(\theta)$ which satisfies $U(\theta, \theta, \phi(\theta)) = U\Big(\theta, \frac{\theta+\theta_{\max}}{2}, \ell_{\max}\Big)$ when the type space is uniformly distributed. Equilibrium 1 is the LME. It defeats equilibrium 2 because, if the belief associated with action $\ell_2^* + \epsilon$ for a small ϵ is of the type whose action is $\ell_2^* + \epsilon$ in equilibrium 1, then that belief cannot sustain equilibrium 2. The proof also shows that the orange equilibrium Pareto dominates the yellow equilibrium by using the indifference condition of type θ_1^* together with the properties of the LME.

3.3 Testable predictions

To highlight the signaling content of parental leave, I consider the consequences a leave extension has on the distribution of leave duration and wages.

Proposition 3 (Relaxing the action space constraint)**.** *When $\ell_{\max}$ is raised:*

1. *(Fully separating) The Pareto optimal and uniquely undefeated fully separating equilibrium (if exists) has every type taking more leave; wages are unchanged.*

2. *(Partial pooling) The partial pooling LME has a smaller fraction of poolers at $\ell_{\max}$ and the previously unconstrained types take longer leave.*

3. *(Pre-pooler post-nonpooler) Signaling has a positive average effect on wages for those who would have pooled before the change and not after the change.*

4. *(Pre-pooler post-pooler) Signaling has a negative effect on wages for those who would have pooled before the change and also after the change.*

The proof can be found in Appendix B.3.

The first two predictions highlight a distinct impact of the signaling channel on the leave distribution, namely that in equilibrium women take into account the actions of other women when making their leave choice. Whether the starting point is a fully separating equilibrium or a partial pooling one, when the action space constraint is relaxed, the leave distribution changes even for women for whom the initial constraint appeared to be non-binding. Intuitively, women take less leave than first-best to differentiate themselves through signaling, but when the less committed types take more leave, the more committed types can also take longer leave while maintaining the signaling value of their choice.

The last two predictions are direct implications of belief monotonicity of the wage function. Workers who go from pooling to not pooling can now reveal their types above average, moving their wages upwards. Workers who continue to pool are now pooling with a group that is less attached on average, resulting in firms believing them to be of worse types and thus lowering their wages. These two predictions also highlight the distributional consequences of signaling, when the same policy change affects wages of different groups in different directions.

The following corollary provides additional predictions of signaling:

Corollary 4 (Relaxing the action space constraint)**.** *When $\ell_{\max}$ is raised:*

1. *Signaling does not have any impact on the wages of mothers who would not take maximum leave both before and after the change.*

2. *At each amount of leave below the original $\ell_{\max}$, mothers taking that amount of leave after the change earn higher wages than mothers taking the same amount before the change.*

While mothers would take longer leave as a result of the policy change even when their previous choices are non-binding, because their types are perfectly revealed in each case, beliefs of the firms are unchanged and therefore wages are unaffected.

When we compare mothers who take the same amount of leave before and after the policy change, the second part of the corollary is a direct consequence of the shifting of the distribution of leave and the monotonicity of leave as a function of type in equilibrium.

4 Empirical setting and specifications

Section 3 suggests that one way to examine the informational content of parental leave is to analyze the changes in the leave distribution and wages that arise from an exogenous change in the maximum allowed leave.

4.1 Parental leave extensions in 1984 and 2002

There have been two major parental leave extensions in Denmark since the 1980s. I use the sharp introductions of both of these extensions to test the comparative statics with respect to the increase in maximum leave duration. The features of the two extensions instruct the construction of the empirical tests of the signaling role of parental leave.

In both cases, the policies went from the proposal stage to the effective stage within three months, and therefore not predictable by potential mothers who were first affected. This feature affirms that, within some window of the program, the relaxation of the action space constraint discussed in Section 3 is indeed an exogenous event.

The first proposal for the 1984 reform began in October 1983 and the reform passed in December 1983.[21] The reform had two steps, which provided 6 weeks and 10 weeks of shareable parental leave respectively, on top of the existing 14 weeks of post-birth maternity leave, all at full benefit compensation.[22] The first step applied to mothers who had given birth within 14 weeks of 1 July 1984, making the effective cutoff date 25 March 1984. The second step applied to mothers who had given birth within 20 weeks of 1 July 1985, making the effective cutoff date 11 February 1985. For the 1984 reform, I focus on the first policy expansion which potential mothers would not have been able to plan around.

Total maternity and parental leave remained at 24 weeks total until a 2002 reform. In 1994, in addition to maternity and parental leave, mothers were provided with an additional 52 weeks of child care leave; unlike with maternity and parental leave which would have to be taken immediately after childbirth, mothers could use their child care leave quota at any time before the child turned nine at a reduced benefit level of 60 percent benefit compensation. Starting on 1 January 2002, the reform removed child care leave, and at the same time provided an additional 32 weeks of shareable parental leave at full benefit compensation. Because the reform was officially passed on 22 March 2002, mothers whose childbirths fell between 1 January 2002 and 26 March 2002 could opt in to either policy. In total, post-reform mothers were entitled to 46 weeks of paid maternity and parental leave, plus an option to extend to a total of 60 weeks keeping the same total benefit unchanged.

When shareable leave was introduced for the first time in 1984, fathers also became entitled to two weeks of paid paternity leave that they can use in the first 14 weeks after a child is born. The majority of fathers in practice use only this leave portion and not the shared parental leave. Less than 3 percent of fathers use shareable leave in both 1984 and 2002. In 2002, the average total leave a father takes changed from 13 days to 15.4 days with the extension, with minimal shifting in the distribution as shown in Appendix Figure 4. I thus refer to the combined leave period that mothers are entitled to, i.e., maternity leave and shared leave, as parental leave.

[21] Rasmussen (2010) provides a summary of the political process that led to the introduction of the program.

[22] Full benefit compensation provides 90 percent of previous pay.

Figure 2 shows the sharp changes brought about by the two leave extensions. The figure plots the average total parental leave duration together with its 95% confidence interval, for every two-week interval within 12 weeks of the effective policy cutoff dates. The average parental leave duration in the three months leading to the 1984 policy cutoff date is 96 days, whereas after the policy cutoff date, it jumps to 134 days. In 2002, the average parental leave duration changes from 156 days in the three months before the policy change to 276 days in the three months after the policy change. The data appendix in Appendix C explains the construction of the leave duration variable.

On average, while there is expected birth seasonality across different months, as shown in Appendix Figure 2, the samples of parents with children born before and after the policy cutoff dates for both leave extensions are balanced. Appendix Tables 2 and 3 show that, across a variety of characteristics of mothers and fathers, including education, pre-birth age, annual earnings, and years of work experience, there are two instances in which the two samples have significantly different means. In particular, mothers with children born after the 1984 policy cutoff date are somewhat older than their pre-reform counterparts, and fathers with children born after the 2002 policy cutoff date are somewhat younger than their pre-reform counterparts.

In 1982, mothers who gave birth around the 1984 leave extension were around 26 years old, earning approximately 25,000 USD (measured in year 2000), and having worked for 5.4 years. In 2000, mothers who gave birth around the 2002 leave extension were around 29 years old, earning approximately 28,000 USD (measured in year 2000), and having worked for 6.5 years. As expected, the average educational attainment of mothers are higher in 2000 than in 1982, with 36 percent of mothers with further education in 2000 compared to 26 percent in 1982.

4.2 Sample definitions

To test the signaling implication of the leave extensions on wages, I introduce subsamples of employed mothers based on their pooling status. I restrict the data to the set of mothers who give birth between a window of 85 days on each side of the policy cutoff dates and who used parental leave.[23] I define the three groups of interest: pre-pooler post-nonpooler (PN) mothers, who would have pooled at maximum leave before but not after the leave extension; pre-pooler post-pooler (PP) mothers, who would have pooled under both policies; and pre-nonpooler post-nonpooler (NN) mothers, who would have taken less than maximum leave under both policies.

I observe each mother under one policy regime and obtain the total leave duration she takes before coming back to work. To approximate each mother's counterfactual leave choice under the other policy, I match the samples of mothers who gave birth before and after the two leave extensions based on their pre-birth characteristics. This leads to a treatment group consisting of the matched sample of mothers who give birth after the reform and a control group consisting of the matched sample of mothers who give birth before the reform. Among mothers who give birth before the cutoff in 1984 (2002), I refer to those who take exactly 14 weeks (24 weeks) of parental leave as pre-poolers and the rest are pre-nonpoolers. I refer to mothers who give birth after the cutoff in 1984 (2002) and take exactly 20 weeks (46 weeks) of parental leave as post-poolers and the rest are post-nonpoolers.[24] A pre-pooler can either be in the PN group or the PP group depending on how many days of leave is taken by the post-reform mother with whom she is matched. The PN group consists of matched pairs of pre-poolers and post-nonpoolers, while the PP group consists of matched pairs of pre-poolers and post-poolers. Likewise, a post-nonpooler can be in either the PN or the NN group depending on how many days of leave is taken by the pre-reform mother with whom she is matched.

I implement a coarsened-exact match following Iacus et al. (2012) using observed pre-birth characteristics of mothers. I match exactly on mothers' educational groups, fathers' educational groups, marital status, and birth parity.[25] I match coarsely on mother's age, father's age, grandmother's work experience, and income in the year

[23]The 85-day window is chosen because in my data, the leave duration variable is available starting 1 January 1984. See Appendix C for more details on the variables and data construction.

[24]For the 2002 policy, I restrict the sample to those who do not use child care leave as a continuation of parental leave. See Section 5.4 for details.

[25]The educational groups are primary education, high school and vocational education, and further education. Marital status includes single, married, or cohabited.

before childbirth.[26] The match has two purposes. First, it facilitates the construction of mothers who would be predicted to fall into the relevant pooling status group under both the actual policy they face, and the counterfactual policy. Second, it helps maintain a balanced sample to counteract possible seasonality, which can change different parts of the distributions of mothers after the policy change.[27] The matching procedure also provides associated weights for all matched observations, and I use these weights in all subsequent results.[28]

Tables 1 and 2 show the summary statistics for the different sample groups associated with the two leave extensions. In both cases, the PN and NN groups are more similar to each other than to the PP group. Relative to the PP group, the PN and NN groups contain mothers who at the time of childbirth were older, having higher annual earnings, and more educated. This is also the case for the same characteristics of the fathers.

Tables 3 and 4 present information that each of the matched subsamples is indeed balanced. Importantly, the tables show that the subsamples are balanced pre- and post-extension even for characteristics not matched on. These include years of labor market experience for both mothers and fathers, as well as the fraction of mothers who work in the public sector, the fraction of mothers who work full-time, or the average social benefits used by mothers.

4.3 Estimating equations and identification

Given the unanticipated nature of the policy changes in my setting, I compare mothers who give birth before and after the policy change. The following empirical specification allows me to assess the impact of the policies over time, as well as to

[26] As explained by Iacus et al. (2012), these variables are recoded into bins based on a pre-defined algorithm.

[27] The chosen variables include characteristics that are both observed and unobserved by the firms. The main objective is to achieve balance while not compromising sample size due to the curse of dimensionality. I obtain similar results varying the set of matched variables, for example using industry group instead of grandmother's work experience.

[28] Different control units may be matched to the same treated unit or vice versa, and the weights account for the different stratum sizes maintaining balance of the matched sample.

directly evaluate the common trend assumption for the different groups. For each outcome y for individual j at time t, the main estimating equation is as follows:

$$y_{jt} = \sum_{k=1}^{n} \left(\beta_{0k} + \beta_k \mathbf{1}^{j}_{\text{birth-after-cutoff}} \right) \cdot \mathbf{1}_{\{t=t_k\}} + \gamma X_j + \epsilon_{jt_k} \tag{2}$$

Workers are compared at each period in the set $\{t_k\}_1^n$ which depends on the leave extension. The pre-birth individual-level control variables represented by X_j consist of fixed effects for mother's and father's age at birth, mother's and father's education subgroup, mother's county of residence, marital status prior to the birth, and an indicator for the child's gender.[29] To facilitate visualization of the effects over a long period of time, a period after the policy change is 4 years for the 1984 leave extension. In both cases, I allow for the pre-trend comparison to occur every year for up to 5 years before the policy change. This means that for the 1984 policy change, I start in 1980, the year my data start. To account for mechanical differences between mothers who give birth around the policy changes, I further combine the two years 1983 and 1984, as well as 2001 and 2002 in the specifications for the two policies respectively. In the year leading up to the policy change, one set of mothers had already given birth, and in the year of the policy change, some of the mothers with childbirth after the policy took effect were still on leave.[30] While treatment varies at a finer level and is measured every year, I cluster at the individual level to account for potential serial correlations in outcomes across periods.[31]

[29]This is the same set of variables used in a similar setting studying maternity leave extensions in Norway by Dahl et al. (2014). Child care subsidy is provided by the local authority, and county of residence helps compare mothers in the same local labor market sharing similar alternative child care arrangements. The results are qualitatively insensitive to the set of control variables.

[30]Mechanically, pre-reform mothers have lower income in the year before the change, but higher income in the year after the change. These effects do not necessarily net out. I represent them together to maintain the visual representation of the policy impacts beyond the initial years. In particular, for the 1984 leave extension, $t_k \in \{1980, 1981, 1982, 1984, 1988, 1992, 1996, 2000, 2004, 2008, 2012\}$. While details about annual earnings from tax income sheets are available until 2012, the hours and therefore wage measure in my data end in 2010. For these variables, the 2012 group represents only 2 years worth of data. For the 2002 leave extension, $t_k \in \{1997, \ldots, 2000, 2002, \ldots, 2012\}$.

[31]I also add year fixed effects to account for the fact that each period contains several years.

The indicator $\mathbf{1}^{j}_{\text{birth-after-cutoff}_j}$ represents the treatment, i.e. the policy change, with mothers whose childbirths were after the cutoff for longer leave eligibility being compared to those whose childbirths were before the cutoff. The parameter β_k represents the difference in period k outcomes between mothers who give birth before and mothers who give birth after the policy change. This specification allows me to evaluate whether mothers in the control and treatment groups following similar trends before the policy change, and the policy's impact on their subsequent labor market outcomes.

The main identification assumption is that mothers who give birth before and after the policy change are not systematically different in absence of the policy change, for example, due to birth seasonality, and that the two groups within each subsample are indeed counterparts of each other. The matching described in Section 4.2 helps obtain this necessary balance. The sample window chosen in the main specification is 85 days on each side of the cutoff. The window is short enough that the policy announcements were unanticipated by all mothers in my samples but long enough to gain power. To address potential concerns that some selected group of mothers may delay their birth delivery right around the cutoff dates to qualify for longer leave eligibility, I exclude a donut hole of 5 days around the cutoff dates.[32] Further robustness checks are presented in Section 5.4.

5 The signaling role of parental leave

The empirical analysis evaluates the role of parental leave extensions in changing the leave taking distribution and their impact on labor market outcomes due to the *relative* length of leave taken by different subgroups.

5.1 Summary of the predictions

I start by summarizing the predictions of the signaling content of parental leave in Section 3 to be empirically evaluated in the context of a policy change that extends

[32]The results are not sensitive to the choice of the window or donut hole period.

the maximum allowed paid leave duration. The prediction descriptions follow the definitions of the subsamples in Section 4.2.

A pure signaling model predicts the following:

Prediction 1 (Shifting of the leave distribution)**.** The distribution of parental leave distribution should shift upwards for mothers for whom the shorter maximum leave duration before the extension would not have been binding.

Prediction 2 (Composition at each leave amount)**.** The average wage is higher for mothers who take a given amount of leave after compared to before the extension.

Prediction 3 (Pre-pooler post-nonpooler subsample)**.** Mothers who would take maximum leave when the allowed duration is shorter but would take less than the maximum when the allowed duration is extended should gain in wages and income as a result of positive belief updating.

Prediction 4 (Pre-pooler post-pooler subsample)**.** The fraction of poolers at the maximum duration decreases when the maximum allowed leave increases. Consequently, mothers who would take maximum leave both before and after the extension would have lower wages and income as a result of negative belief updating.

Prediction 5 (Pre-nonpooler post-nonpooler subsample)**.** The signaling channel has zero effects on mothers who would not take maximum leave before and after the extension.

Since the predictions above are stated for a pure signaling model, I now consider how the presence of other channels might affect these predictions. If there were direct effects of leave taking on productivity, then we would still be able to evaluate whether signaling plays a role by examining predictions on the leave-taking distribution as well as predictions on wages and income. Such channels would not account for predictions that pertain to how one's behavior is influenced by that of others. In particular, Predictions 1 and 2 arise through the informational content of parental leave and therefore would indicate the relevance of signaling even in the presence of other channels. A direct effect of leave taking on productivity might alter the level of income and wages in Predictions 3 and 4 but would not affect the relative comparison between the PP and PN groups. In particular, if the PN group has higher wages after a leave extension, then signaling would predict that the PP group has even higher wages, and if the PP group has lower wages after a leave extension, then signaling would predict that the PN group has even lower wages. Section 5.5 further the discussion of other channels and their potential contribution to the patterns seen in the data.

5.2 Prediction on leave-taking distribution

Prediction 1

Consistent with Prediction 1, the leave distribution of unconstrained mothers shifts up. I evaluate this prediction by examining how the empirical cumulative distribution of parental leave changes when the maximum allowed leave duration is extended. Figure 3(a) shows the distribution of leave taking before and after the 1984 policy change. While the fraction of mothers taking less than 12 weeks is similar, the mothers who give birth post-reform are substantially less likely to choose a leave duration within two weeks of the pre-reform maximum.[33] Despite the low level of leave-taking and correspondingly high fraction of pooling at the maximum allowed leave duration in 1984, the evidence is qualitatively consistent with Prediction 1.[34]

Measurement error in leave taking presents a potential threat to this interpretation. If some fraction of poolers are recorded as returning to work within one week of the maximum, then we would indeed observe an upward shift in the distribution of leave taking after the policy change concentrated at the high end of the leave distribution. However, in that case, we would expect to find that mothers who pool at the maximum do not systematically differ from those who take a few days less leave. Appendix Figure 3 demonstrates that the differences in leave taking translate into meaningful changes in wages. Mothers who take the maximum allowed leave

[33] A fraction of about 5 percent of mothers take between 1 and 12 weeks under both policies, which could reflect that only mothers who have a fixed preference for leave taking would take leave at such a low level.

[34] The fraction of poolers goes from about 88 percent to 77 percent after the 1984 policy change.

duration earn about 2.5 percent percent higher wages before childbirth and up to 8 percent higher wages 20 years after compared to those who take within 10 days of the maximum. This supports the interpretation that the shift in the distribution of leave taking after the 1984 policy change represents an interdependency in choices based on the information that a mother's choice conveys to the firm as opposed to error in measuring leave taking.

Turning to the 2002 policy change, Figure 3(b) offers clear evidence of an upward shift throughout the distribution of leave taking. Consistent with Prediction 1, the shift in the distribution suggests that mothers for whom the maximum allowed leave duration was not binding choose to take more leave after the reform. As with the 1984 result, the results are consistent with mothers choosing a higher leave duration as a consequence of others' choices rather than a direct effect of the policy. A slight complication arises when considering the 2002 leave extension due to the option to take an additional 52 weeks of child care leave. Consider a mother who takes less than the maximum allowed 24 weeks of maternity leave at full benefit and also takes some child care leave. She might choose to take additional parental leave at the full benefit level when the reform eliminates the option to take child care leave. To address this concern, Appendix Figure 5 restricts the post-reform mothers to those whose children were born between 1 January 2002 and 27 March 2002 when mothers could choose whether to be compensated under the pre-reform policy or the post-reform leave policy.

The pre-nonpooler post-nonpooler (NN) group under each leave extension gives another indication of the magnitude of this shift in the leave distribution. Under the 1984 and 2002 extensions, the matched sample of mothers for whom the pre-reform maximum would not have been binding increases leave taking from 81 to 97 days and from 132 to 186 days, respectively.

Prediction 2

This prediction evaluates the composition of mothers who take the same amount of leave before and after the policy change. Table 5 documents the wage differences between mothers who choose to take at most 14 weeks of parental leave before or after the policy change in 1984 as well as mothers who choose to take at most 24 weeks of parental leave (without the use of child care leave in the subsequent 3 weeks) before or after the policy change in 2002.

Columns (1) reports that the group of mothers who choose to take at most 14 weeks of parental leave after the 1984 policy change has higher wages over time than the group of mothers who choose the same level of leave before the policy change. Column (2) adds fixed effects for the number of days of leave and finds a similar pattern, while columns (3) and (4) present analogous specifications for the 2002 policy change. Consistent with Prediction 2, the composition of mothers taking the same amount of leave changes: mothers who choose to take shorter leave when a longer amount of leave is available earn higher wages. The average difference for the 1984 leave extension is about 10 percent across the periods, whereas the average difference for the 2002 leave extension is about 5 percent.

5.3 Predictions on earnings and wages

This section documents that mothers who are able to distinguish themselves by taking *relatively* less leave as a consequence of a leave extension policy gain in wages and income, consistent with the role of belief updating in wage setting. The 1984 policy extending the maximum paid leave duration with no other complementary leave scheme offers a direct test of the theoretical predictions. Consistent with the theory, the fraction of poolers at the longer maximum leave duration after the change is smaller, going from 88 percent to 77 percent. I evaluate both the average impact of the policy change for all mothers, as well of the pre-pooler post-nonpooler (PN) and pre-pooler post-pooler (PP) groups to evaluate the distinctive feature of the signaling channel which adds to the average effect in opposite directions.

Figure 4 documents the dynamics of the treatment effect on wages and income of the 1984 policy change. The figure shows that mothers whose children were born before and after the policy change are similar and were following common trends before their divergence due to one group being exposed to the leave extension and the other not. On average, the policy does not have any effect on hourly wages or annual earnings when comparing all mothers whose childbirths are after versus before the policy change. The estimates are precise, ruling out a difference in hourly wages greater than 1 percent and a difference in annual earnings greater than 750 USD.[35] The average effects are consistent with the findings by Rasmussen (2010) studying the same policy change, as well as various studies of other leave extensions in OECD countries.

In contrast to the average effect, the following discusses how the leave extension has a positive impact on wages for the PN group but a negative impact for the PP group, consistent with the signaling role of parental leave choice.

Prediction 3

Figure 4 shows that mothers in the PN group earn 2 percent to 5 percent more in the years following the policy change. Consistent with the theoretical prediction, the extension enables these mothers to distinguish themselves by taking less leave relative to the maximum duration. These mothers experience positive effects on earnings as a consequence of the policy change even though they take more leave (an average of about 16 days) in absolute terms. The effects are persistent, lasting throughout the 20-year period that we observe the mothers in the labor market after childbirth. When considering annual earnings, which include observations when a mother is not employed, the birth-after-cutoff effects for the PN group ranges between 1,500 USD and 3,000 USD, or about 5 percent to 10 percent relative to the base of 30,000 USD. The relatively larger effect for income compared to wages is consistent with PN mothers working more hours if they give birth after the policy change. The point estimates in Appendix Figure 6 and Figure 10 indeed show a difference in log annual work hours and employment status for the PN group. While the theoretical model abstracts from the choice of work hours within a period, the positive effect on work hours could be consistent with an endogenous response of hours to wage increases or with a model in which firms change wages by assigning the workers to positions with

[35] Income variables reported are all adjusted to index to year 2010.

more intensive hours.

Prediction 4

For the PP group, the policy change has a negative effect on wages of around 1 percent to 2 percent, and a negative effect on annual earnings of around 250 USD to 1,000 USD. The magnitude of the change for the PP group is smaller than that of the PN group. Since the fraction of poolers decreases by about 12 percent after the 1984 policy change, the relatively small change for the PP group is consistent with the leave extension resulting in a smaller change in firms' beliefs about mothers who take the maximum leave duration.

Prediction 5

Figure 6 shows that the effects of the leave extension on wages and income for the NN group are small and insignificant. While they are less precisely estimated due to the smaller sample size, the point estimates indeed fall between those of the PN and PP groups, consistent with Prediction 5. Given that the characteristics of the NN group are more similar to those of the PN group than the PP group as shown in Table 1, and the average change in leave duration as a result of the policy change is around 16 days for both NN and PN, the findings are difficult to reconcile with these observed differences alone. Section 5.5 further discusses other channels through which parental leave extension can affect workers.

Discussion of magnitudes

The findings in Figure 4 suggest that signaling can have long-lasting labor market consequences. The persistent effects suggest that employer learning, as outlined in Farber and Gibbons (1996) and Altonji and Pierret (2001), does not outweigh the direct impact of being perceived as a better worker during one's prime working age. A better position at the current firm can lead to better subsequent training and efficient human capital acquisition or higher bargaining power in labor markets with asymmetric information across employers, consistent with the findings in Kahn (2013); Kahn and Lange (2014) and the framework from Bernhardt (1995).[36]

An important caveat in interpreting the magnitudes is that the difference between the PN and PP groups does not represent a predicted increase in wages from choosing not to take the maximum allowed leave duration. First, the positive effect on wages for this group does not capture an effect of taking less parental leave in absolute terms. Second, since the choice of leave duration is an equilibrium outcome, the PN group consists of mothers for whom such wage gains from taking relatively less leave would be possible. The design of the empirical tests allows us to define such groups to evaluate the relevance of signaling. Thus, while the PN group on average takes 16 more days of leave as a result of the 42-day extension, the results do not imply that taking 28 days less than the maximum allowed duration would result in the same wage gain for any particular individual.

5.4 Robustness of main results

Additional test using 2002 policy change

The 2002 policy change offers a complementary test of Predictions 3 and 4 for a separate group of workers under different labor market conditions. The existence of the child care leave scheme prior to 2002 complicates constructing the direct empirical counterparts of the comparative statics. I thus consider a modified definition of pooling in the pre-reform period for this alternative test, omitting mothers who came back to work after using child care leave as a continuation to parental leave. I define pre-poolers as mothers in the pre-reform period who come back to work after exactly 24 weeks of full-benefit leave without using child care leave. Similarly pre-nonpoolers are mothers in the pre-reform period who come back to work before 24 weeks of full-benefit leave without using child care leave. Mothers who take at least 46 weeks of leave in the post-reform period form the post-pooler group, and the rest

[36]Within the same firm, Baker et al. (1994) and Gibbons and Waldman (1999) show that workers who receive a large wage increase early in a ladder in their career are those who are quickly promoted to the next level.

are post-nonpoolers. I then define the PN, PP, and NN groups as in Section 4.2. Even with this restriction, the simplified model in Section 3 does not make a prediction about the direction of changes in beliefs about either the PN or PP groups, which would depend on which types of mothers pool at 24 weeks and forgo child care leave in equilibrium before the policy change.

Although it is theoretically ambiguous how firms' beliefs would change given the child care option before the leave extension, the signaling model still makes a prediction about the PN group relative to the PP group. In particular, the policy change induces separation among the pre-pooler mothers, and in any equilibrium in which beliefs are monotonic in leave duration after the reform, the policy should have a more positive impact on wages and income for the PN group relative to the PP group. Thus we have the following modified version of Predictions 3 and 4 for the 2002 policy change.

Prediction 6 (PN compared to PP)**.** The effect of the 2002 leave extension should be relatively more positive for the PN group compared to the PP group.

Figure 5 shows qualitatively similar results using the 2002 leave extension. As with the 1984 leave extension, the average effects are small and insignificant. The 2002 policy change has a positive effect on income and wages for the PN group, but a negative effect for the PP group, with the difference being 2 percent to 5 percent in hourly wages and about 2,500 USD to 3,700 USD in annual earnings.[37] Figure 6 confirms that the effects for NN group lie in between those for PN and PP.

Falsification test

Figure 7 presents the results of a falsification exercise which demonstrates that the main results do not arise due to the sample selection procedure. For each policy change, I define a placebo cutoff date three months after the date when the old policy

[37]The positive wage effects in 2002 and 2003 include mechanical effects due to how the data are recorded. In the data, income includes paid leave but the hours measure does not consist of the leave period. , Due to the long leave period in 2002, by definition, more mothers in the PP group are still on leave at the end of 2003, deflating their measured hours as seen in Appendix Figure 6 and inflating the estimated wage effects.

was no longer in effect.[38] In both cases, I apply the same procedure in Section 4.2, matching mothers with children born in the three-month windows around the placebo dates to mothers with children born before the actual cutoff dates, and use the same estimating equations as in Section 4.3. Each placebo policy change shows a precisely estimated null result for both the PP and PN groups.

5.5 Evaluation of non-signaling channels

Productivity effect of leave extension

The evidence in Figure 6 suggests a limited role for a direct effect of days of leave in accounting for wage changes following the 1984 and 2002 extensions. The NN group increases leave taking from 81 to 97 days in the 1984 sample and from 132 to 186 days in the 2002 sample. In both cases, consistent with a pure signaling model (see Prediction 5), the point estimates for the effect of the policies on wages and annual earnings are close to zero. While productivity effects due to leave taking may arise at longer periods of absence, I find little evidence for this channel within the range of leave observed in my data.[39]

Analyzing the relationship between days of leave and wages at different levels of leave-taking provides further evidence to examine the potential role of productivity effects. Following the 1984 policy change, while wages do not change for the NN group when leave increases on average from 81 to 97 days, the PN group experiences wage gains when leave increases on average from 98 days to 114 days, and the PP group experiences wage losses when leave increases on average from 98 days to 140 days. The 2002 policy change shows similar patterns, with no wage change for the NN group increasing leave on average from 132 to 186 days, positive wage changes for the PN group increasing leave on average from 168 to 220 days, and negative wage

[38]The 1984 placebo cutoff date is three months after 25 March 1984, and the 2002 placebo cutoff date is three months after 27 March 2002. Between 1 January 2002 and 27 March 2002, mothers could opt into either the old or new policy, so the placebo cutoff date avoids having the estimates being influenced by this change. In practice, the majority of mothers elected to use the new policy.

[39]In a cross-country study, Ruhm (1998) finds that longer leave duration starts to have a negative effect on wages after 9 months.

changes for the PP group increasing leave on average from 168 to at least 322 days. Although there are compositional differences between these groups, an explanation based solely on productivity effects would not predict this non-monotonic relationship between wages and days of leave.

Differences in observable characteristics

The theoretical framework delivers distinct predictions for mothers based on their pooling decision. To determine whether observable differences across the groups directly contribute to the differences in their outcomes, Figures 8 and 9 examine the heterogeneous effects of the policy changes. For each leave extension, I estimate equation (2) separately on log hourly wages for mothers giving birth within the three-month window around the cutoff date for each of the following groups: above and below the median age at childbirth, below and above the median work experience by the time of the policy change, by three educational groups, and by pre-birth annual incomes below or above median. Although analyzing groups of mothers based on pooling behavior reveals substantial differences in the effect of the two reforms, separating the sample of mothers instead based on the observable characteristics listed above yields small and insignificant differences. When there are differences between subgroups, the pattern is not robust across the two policies, and the magnitudes cannot account for the differences between the PN and PP groups.

Negative selection into employment

When interpreting the negative effect on wages for the PP group as evidence of the role of signaling in determining firms' beliefs and consequently wages, a potential concern may arise due to selection into employment. In particular, a leave extension can make it easier for women to combine work and family, and women who are less attached to work may then be more likely to remain in the labor force.[40] Figure 10

[40]Blau and Kahn (2013) and Thomas (2018) suggest that family-friendly policies can result in statistical discrimination against women in upper-level positions due to this negative selection. Blau and Kahn (2013) document that non-US OECD countries have higher female labor force participation, but lower rates of women working full-time and in managerial and professional occupations.

shows evidence that negative selection into employment does not pose a concern. I find no effect on employment for the PP group under either of the leave extensions. In fact, the PN group exhibits a positive effect on wages despite an increase in employment. This further suggests a small role for negative selection, which might be overwhelmed by channels that lead to positive long-term effects such as better subsequent training.

Learning and spillovers

I assess the plausibility of explanations that rely on (1) mothers resolving uncertainty about the employment costs of motherhood as they spend more time away from the labor force, and (2) mothers learning from one another's leave taking behavior. Under this view, when the policy change relaxes the constraint on the maximum allowed leave duration, mothers who would choose to pool at the maximum before the leave extension would increase their choice of leave. To the extent that the additional time away from the labor force leads them to resolve additional uncertainty about the value of staying at home relative to returning to work, these mothers may separate into two groups: those with weaker preferences for work would take the new maximum leave duration and form the PP group, earning relatively lower wages; those with stronger preferences for work would take fewer additional weeks of leave and form the PN group, earning relatively higher wages. A ripple effect, whereby mothers who would take less than the maximum allowed leave duration also change their behavior in response to the leave extension, can arise if mothers observe their coworkers' decisions and follow them, even if leave choices do not convey any information to employers. Several results point against this combination of factors as an alternative explanation for the patterns in the data.

First, for each of the reforms, wages and earnings do not significantly differ between the PN and PP groups in any of the years prior to the reform, and the placebo policies (Figure 7) also do not show any differences between PN and PP. These results pose challenges for explanations that rely on differences between the two groups in the strength of their preferences for work, as the learning process that separates the groups would have to do so only if they take at least 14 weeks of leave.

Second, as Figure 2 shows, the increase in leave taking occurs sharply after the policy change and then remains stable. This pattern does not provide support for the hypothesis that the ripple effect arises due to mothers observing coworkers' decisions, which would predict a gradual increase over time.

Third, Table 6 shows a no effect on future leave taking and fertility decisions, which further suggests a limited role if any for learning-based explanations. I evaluate the impact of the two leave extensions using following estimating equation:

$$y_j = \gamma_0 + \gamma_1 \mathbf{1}^j_{\text{birth-after-cutoff}} + \gamma X_j + \epsilon_j \tag{3}$$

The outcome variables of interest for each individual j are the total number of children by year 2012, the distance between the birth of the current child and the next child (if any), and parental leave duration for the next child (if any). I use the same control variables X_j as in Section 4.3. The results show a negligible impact of the reforms on fertility and subsequent leave taking decisions.[41]

6 Concluding remarks

This paper introduces a model of parental leave as a signal of labor market attachment for mothers, derives testable predictions through comparative statics when the maximum allowed leave increases, and tests these predictions using administrative longitudinal data from Denmark and exogenous policy changes. More generally, the results in the paper suggest that there can be unintended consequences when workers' take-up of parental leave or other benefit programs conveys information to firms. In the particular case of paid leave policies, workers take less leave than they would

[41]The findings are consistent with the cross-country analysis by Olivetti and Petrongolo (2017) and the study of seven expansions in Norway by Dahl et al. (2016). One exception is a study by Lalive and Zweimüller (2009), which finds that expanding leave duration from one year to two years in Austria in 1990 significantly increased the probability of having a second child for first-time mothers. In their setting, mothers who have another child while within 3.5 months of the previous leave period would gain an automatic renewal, which helps explain the effect.

in the absence of this channel, and a leave extension reveals additional information about their types which causes some workers to gain in wages at the cost of others.

Paid parental leave in the United States is presently provided at the firm level. Since workers have private information, firms may use benefit programs as screening device (Stiglitz, 1975; Rothchild and Stiglitz, 1976). Future work can explore how the signaling channel changes our understanding of the potential effectiveness of these arrangements.

Signaling also provides a justification for mandatory minimum parental leave policies. The Danish government implemented a two-week minimum post-birth maternity leave starting in the 1990s. While mandatory minimum leave may confer benefits to children and the government puts a higher welfare weight on those benefits than mothers do, reducing wasteful signaling provides a complementary explanation. Analyzing alternative leave policies under asymmetric information would provide an important step towards designing optimal policies.

The signaling channel provides a new perspective on why fathers do not take up the shared leave even when mothers do not fully exhaust the available benefits.[42] It could be the case that employers' beliefs are strongly influenced by men's leave taking, which in turn reinforces itself by preventing men from deviating and taking up more of this form of child care duties.[43] This is consistent with a model of social norm as a signaling equilibrium (Bernheim, 1994).[44] Future work can examine the role of signaling and gender differences in other areas of the labor market.

[42] This is the case not just in Denmark but in other countries with shared leave. See summary by Adema et al. (2016).

[43] Haas et al. (2002) and Bygren and Duvander (2006) survey employees in firms in Sweden and find that employer attitudes and workplace culture are important determinants of fathers' take-up of shared leave.

[44] Alternative explanations such as within-household division of labor (Becker, 1985) may be more important for explaining gender differences in the child rearing process compared to a one-time and relatively short-term occurrence such as parental leave.

A Parental-leave programs in Denmark

The history of parental leave programs in Denmark goes back over 125 years (OECD, 2017).[45] The government guarantees a standard level of paid leave by reimbursing firms directly, and the total leave length has fluctuated anywhere between 14 weeks of post-birth leave in the early days to a year or more in recent times.

Between 1892 and 1965, only women working in factories were eligible for maternity leave, ranging from two weeks to 14 weeks, with paid leave being introduced in 1933. Starting in the 1980s, this occupation restriction was fully lifted, making paid leave widely available as long as mothers were eligible for the jobseeker allowance, which covers the majority of workers and active unemployed workers. The first major change occurred in two steps in 1984 and 1985, expanding post-birth maternity leave from 14 weeks to add an additional six weeks of shared parental leave in in 1984 and ten weeks in 1985. Starting at the same time, fathers were entitled to two weeks of compensated leave within 14 weeks of the arrival of a child. Fathers and mothers cannot take shared parental leave at the same time, and in practice, the incidence of parental leave take-up falls mostly on mothers. Maternity and parental leave is effectively paid at a wage replacement rate of 90 percent. The Directive on Equal Treatment of Men and Women in 1989 formally established job protection for women during maternity leave, as well as a minimum requirement of a two-week absence for mothers after birth.

The 1990s witnessed an introduction to a child care leave scheme for women with a child below the age of nine, which provided additional paid leave, and, depending on the exact period, with the wage replacement rate ranging between 60 percent and 90 percent for up to a period of 52 weeks. The rationale for this program was in part to address high unemployment rates by encouraging rotation of the workforce when unemployed individuals can fill in temporary positions to gain work experience (Westergaard-Nielsen, 2002)

The last major change up to 2012 was a reform in 2002 which once again simplified

[45]The OECD report summarizes information from Borchorst (2006), Rasmussen (2010), Pylkkänen and Smith (2004), and Nielsen et al. (2004) among others.

the system and abolished child care leave. Instead, the reform extended post-birth shared parental leave to 32 weeks which can be further extended to 46 weeks with the same amount of total pay, on top of the existing 14 weeks of maternity leave.

Among OECD countries, Denmark is at the median with regard to total allowed leave duration at about one year of paid leave, a similar level to Canada and Poland, and slightly below that of Sweden and Germany. Throughout the period that Denmark provides paid parental leave to its mothers, the country also has a universal system of generous and high quality child care.[46] The Danish setting represents an acknowledged standard for parental leave and allows an analysis that focuses on the impact of leave on female labor supply.

B Theoretical model

B.1 Proof of Proposition 1

(1) Characterization

Mailath (1987) provides a set of sufficient conditions for the existence of a fully separating equilibrium in a setting with an unbounded action space. We apply the framework of Mailath and von Thadden (2013) because the action space is bounded. It suffices to verify that the following conditions hold.

1. Smoothness: $U(\theta, \hat{\theta}, \ell)$ is C^2 on $[\theta_{\min}, \theta_{\max}]^2 \times [0, \ell_{\max}]$

2. Belief monotonicity: U_2 never equals zero

3. Type monotonicity: U_{13} never equals zero

[46]According to The Ministry of Social Affairs in consultation with the Ministry of Education (2000) report, daycare facilities in Denmark in 1999 cover 28 percent of six-month olds and under, 68 percent of one-year olds, and 80 percent of two-year olds. As summarized by Datta Gupta and Simonsen (2010), daycare is of exceptionally high quality in Denmark, with an average staff-to-child ratio of 1-to-7 and extensive qualification requirements for staff. Parents pay a maximum of 33 percent of the total costs, which equates to an upper bound of $3,000 in child care costs per year.

4. The first-best contracting problem (the problem under full information), $\max_{\ell \in [0,\ell_{\max}]} U(\theta, \theta, \ell)$, has a unique solution ℓ^* for all $\theta \in [\theta_{\min}, \theta_{\max}]$.

5. (i) For all $\theta \in (\theta_{\min}, \theta_{\max})$, $U_{33}(\theta, \theta, \ell^*) < 0$. (ii) There exists $k > 0$ such that for all $(\theta, \ell) \in [\theta_{\min}, \theta_{\max}] \times [0, \ell_{\max}]$, $U_{33}(\theta, \theta, \ell) \geq 0$ implies $|U_3(\theta, \theta, \ell)| > k$.

Conditions 4 and 5 adapt the local concavity conditions of Mailath (1987) to a setting in which the action space is unbounded.

- Given $U(\theta, \hat{\theta}, \ell) = (y - \tau) + \tau\Big(1 - F_\epsilon\Big(\hat{\theta}(\ell)C(1) - (y - \tau)\Big)\Big) - \theta_i C(1 - \ell)$, the smoothness condition is satisifed as long as the density of ϵ is continuously differentiable.

- Since $U_2 = -\tau C(1) f_\epsilon\Big(\hat{\theta}(\ell)C(1) - (y - \tau)\Big)$, belief monotonicity holds as long as ϵ has full support on $[\theta_{\min}, \theta_{\max}]$.

- Since $U_{13} = C'(1 - \ell)$, type monotonicity holds as long as $C(\cdot)$ is strictly monotone.

- Since $U_3 = \theta C'(1 - \ell)$, the unique solution to the first-best contracting problem is given by $\ell^* = \ell_{\max}$.

- Since $U_{33} = -\theta C''(1 - \ell)$, condition 5(i) holds as long as $C(\cdot)$ is convex. Convexity $C(\cdot)$ also ensures that condition 5(ii) holds vacuously.

Now the solution to the differential equation is given by

$$\begin{aligned} \frac{d\ell}{d\theta} &= -\frac{U_2}{U_3} \\ &= \frac{\tau f_\epsilon(\theta C(1) - (y - \tau))C(1)}{\theta C'(1 - \ell)}. \end{aligned}$$

Let ϵ be uniformly distributed on $(-(y - \tau - \theta_{\min} C(1), -(y - \tau - \theta_{\max} C(1))$, which is the smallest support such that all types have a positive probability of choosing to work or to not work. Then we obtain

$$\int C'(1-\ell)\,d\ell = \int \frac{\tau}{\theta}\frac{1}{\theta_{\max}-\theta_{\min}}\,d\theta$$
$$-C(1-\ell) = \frac{\tau\log\theta}{\theta_{\max}-\theta_{\min}} + K$$

for any $K \in \mathbb{R}$. Solving for ℓ gives the desired characterization $\ell(\theta) = 1 - C^{-1}\Big(-\frac{\tau\log\theta}{\theta_{\max}-\theta_{\min}} + K\Big)$.

(2) Existence

For existence, we must have $\ell(\theta) \in [0, \ell_{\max}]$ for all $\theta \in [\theta_{\min}, \theta_{\max}]$. Note that $\ell(\theta_{\min}) = 0$ if $K = C(1)+\frac{\tau\log\theta_{\min}}{\theta_{\max}-\theta_{\min}}$, in which case we have $\ell(\theta_{\max}) = 1-C^{-1}\Big(C(1) - \tau\frac{\log\theta_{\max}-\log\theta_{\min}}{\theta_{\max}-\theta_{\min}}\Big)$. Thus, a fully separating equilibrium exists if and only if $\ell(\theta_{\max}) \geq 1-C^{-1}\Big(C(1) - \tau\frac{\log\theta_{\max}-\log\theta_{\min}}{\theta_{\max}-\theta_{\min}}\Big) := \ell_{\text{bound}}$.

B.2 Proof of Proposition 2

(1) Characterization

The following characterizes the class of partial pooling equilibria in which types below some threshold $\theta \in (\theta_{\min}, \theta_{\max})$ separate and types above θ pool at $\ell_{\max}$. For θ to be the cutoff, type θ is indifferent betweeen (i) taking maximum leave and be believed to be the expected type among the poolers, and (ii) taking some amount $\phi(\theta)$ of leave to separate and be correctly believed to be type θ:

$$U(\theta,\theta,\phi(\theta)) = U\bigg(\theta, \frac{\theta+\theta_{\max}}{2}, \ell_{\max}\bigg).$$

Since

$$u(\theta,\theta,\phi(\theta)) = y-\tau+\tau[1-F_\epsilon(\theta C(1)-(y-\tau))] - \theta C(1-\phi(\theta))$$

and

$$u\left(\theta, \frac{\theta+\theta_{\max}}{2}, \ell_{\max}\right) = y - \tau + \tau\left[1 - F_\epsilon\left(\frac{\theta+\theta_{\max}}{2}C(1) - (y-\tau)\right)\right] - \theta C(1-\ell_{\max}),$$

we have

$$\theta[C(1-\phi(\theta)) - C(1-\ell_{\max})] = \tau\left[F_\epsilon\left(\frac{\theta+\theta_{\max}}{2}C(1) - (y-\tau)\right) - F_\epsilon(\theta C(1) - (y-\tau))\right].$$

This implies

$$\begin{aligned}\phi(\theta) &= 1 - C^{-1}\left(C(1-\ell_{\max}) + \frac{\tau}{\theta}\left[F_\epsilon\left(\frac{\theta+\theta_{\max}}{2}C(1) - (y-\tau)\right) - F_\epsilon(\theta C(1) - (y-\tau))\right]\right)\\ &= 1 - C^{-1}\left(C(1-\ell_{\max}) + \frac{\tau}{\theta}\frac{\theta_{\max}-\theta}{\theta_{\max}-\theta_{\min}}\right)\end{aligned}$$

for $\epsilon \sim \mathcal{U}(-(y-\tau-\theta_{\min}C(1), -(y-\tau-\theta_{\max}C(1))$. Note that the degenerate case of a partial pooling equilibrium would be the separating equilibrium that maximizes leave taking since $\phi(\theta_{\max}) = \ell_{\max}$.

Let $\bar{K}$ satisfy $\ell(\theta) = \phi(\theta) = 1 - C^{-1}\left(-\frac{\tau \log\theta}{\theta_{\max}-\theta_{\min}} + \bar{K}\right)$. To satisfy the condition for separation of the types $(\theta_{\min}, \theta)$, we must have $\ell(\theta_{\min}) = 1 - C^{-1}\left(-\frac{\tau\log\theta_{\min}}{\theta_{\max}-\theta_{\min}} + \bar{K}\right) \geq 0$.

(2) Refinement

The characterization above shows how to obtain the class of partial pooling equilibria indexed by $\bar{\theta}$, with types above $\bar{\theta}$ pooling and types below $\bar{\theta}$ separating. The function $\phi(\bar{\theta})$ specifies the separating action chosen by the type $\bar{\theta}$ who is indifferent between separating and pooling. The lexicographically maximum equilibrium (LME) can be obtained by determining the equilibrium that is most preferred by the best type, i.e., $\theta_{\min}$.

In the class of partial pooling equilibria characterized above, the most preferred equilibrium of $\theta_{\min}$ is the equilibrium in which $\theta_{\min}$ is able to take the longest leave; this is because forgoing leave is costly, and their type is correctely inferred to be $\theta_{\min}$ in this class of equilibria because the best types are separating.

To prove that this is the LME, I will consider alternative equilibria classified based on whether $\theta_{\min}$ separates or pools. The following six groups are exhaustive:

1. Pooling on $(\theta_{\min}, \theta_1)$, separating on (θ_1, θ_2)
2. Pooling on $(\theta_{\min}, \theta_1)$, pooling on $(\theta_1, \theta_{\max})$
3. Pooling on $(\theta_{\min}, \theta_1)$, pooling on (θ_1, θ_2), pooling on (θ_2, θ_3)
4. Pooling on $(\theta_{\min}, \theta_1)$, pooling on (θ_1, θ_2), separating on (θ_2, θ_3)
5. Separating on $(\theta_{\min}, \theta_1)$, separating on (θ_1, θ_2)
6. Separating on $(\theta_{\min}, \theta_1)$, pooling on (θ_1, θ_2)

- For any equilibrium in the first group, there is an equilibrium involving separation that the type $\theta_{\min}$ prefers. To see this, let $\alpha(\theta_1)$ denote the amount of additional leave that leaves type θ_1 indifferent between separating and pooling. Note that α is decreasing in θ since additional leave is more valuable for types with higher disutility of effort θ. This implies that decreasing θ_1 results in higher leave taking, and in fact taking $\theta_1 \to \theta_{\min}$ results in the $\theta_{\min}$ type taking longer leave. The $\theta_{\min}$ type prefers this alternate equilibrium because she takes longer leave and is believed to be a better type.

- An equilibrium in the second group satisfies

$$U\left(\theta_1, \frac{\theta_1 + \theta_{\min}}{2}, \ell_{\max} - \eta\right) = U\left(\theta_1, \frac{\theta_1 + \theta_{\max}}{2}, \ell_{\max}\right).$$

Since

$$U\left(\theta_1, \frac{\theta_1 + \theta_{\min}}{2}, \ell_{\max} - \eta\right) = y - \tau + \tau\left[1 - F_\epsilon\left(\frac{\theta_1 + \theta_{\min}}{2} C(1) - (y - \tau)\right)\right] - \theta C(1 - \ell_{\max} + \eta)$$

and

$$U\left(\theta_1, \frac{\theta_1 + \theta_{\max}}{2}, \ell_{\max}\right) = y - \tau + \tau\left[1 - F_\epsilon\left(\frac{\theta_1 + \theta_{\max}}{2} C(1) - (y - \tau)\right)\right] - \theta C(1 - \ell_{\max}),$$

we have

$$\tau\left[F_\epsilon\left(\frac{\theta_1+\theta_{\max}}{2}C(1)-(y-\tau)\right)-F_\epsilon\left(\frac{\theta_1+\theta_{\min}}{2}C(1)-(y-\tau)\right)\right]= \theta[C(1-\ell_{\max}+\eta)-C(1-\ell_{\max})].$$

With ϵ uniformly distributed, this becomes

$$\eta(\theta)=\ell_{\max}-1+C^{-1}\left(C(1-\ell_{\max})+\frac{\tau}{2\theta}\right).$$

Maximizing $U\left(\theta_{\min},\frac{\theta_{\min}+\theta_1}{2},\ell_{\max}-\eta(\theta_1)\right)$ over θ_1 gives

$$\frac{dU}{d\theta_1}=-\frac{\tau}{2}\frac{1}{\theta_{\max}-\theta_{\min}}-C(1-\ell_{\max})<0$$

which implies that U is maximized at $\theta_1=\theta_{\min}$. Thus the type $\theta_{\min}$ prefers a fully pooling equilibrium to an equilibrium in the second group. Under the condition that $\theta_{\min}$ prefers a fully separating equilibrium to a fully pooling equilibrium, the LME cannot be in this group.

- For any equilibrium in the third group, there is an equilibrium in the second group that the type $\theta_{\min}$ prefers. Specifically, consider the alternative equilibrium in which types (θ_1,θ_2) pool at the same level of leave as (θ_2,θ_3). For θ_1 to be indifferent between pooling with (θ_1,θ_3) or pooling with $(\theta_{\min},\theta_1)$, the latter has to be at a higher level of leave since the former is at a higher level of leave, and thus the type $\theta_{\min}$ prefers the alternative equilibrium.

- The argument for the fourth group is analogous to the argument for the first group.

- To have an equilibrium in the fifth group, it must be the case that the DE is satisfied at θ_1. Thus, we are either in the case of a fully separating equilibrium, or we are in the case of the sixth group. As noted previously, the class of partial pooling equilibria characterized above includes the fully separating equilibrium as a special case, so this cannot be more preferred to the candidate LME which is the partial pooling equilibrium in which $\theta_{\min}$ takes the longest leave.

- Since the sixth group includes the class of partial pooling equilibria characterized above, assume $\theta_2 \neq \theta_{\max}$. Then the argument is analogous to the argument for the first group or the argument for the second group.

Applying Theorem 1 in Mailath et al. (1993), the LME is undefeated.[47]

(3) Property of the refinement

Denote the LME by the type θ_1^* who is indifferent between separating and choosing $\phi(\theta^*)$ or pooling at $\ell_{\max}$. Consider a partial pooling equilibrium $P(\theta_1)$ with $\theta_1 < \theta^*$ as the indifferent type. Since $\phi(\theta)$ is increasing, the action $\phi(\theta_1) + \epsilon$ with $\epsilon \to 0$ is off the equilibrium path for $P(\theta_1)$ but not for $P(\theta^*)$.

The belief upon observing action $\phi(\theta_1) + \epsilon$ that the individual's type is consistent with the type who chooses $\phi(\theta_1)+\epsilon$ on path in $P(\theta_1^*)$ cannot be sustained in equilibrium; type θ_1 would prefer to deviate and take ϵ greater leave to be believed to be a better type.

For a partial-pooling equilibrium $P(\theta_2)$ with $\theta_2 > \theta^*$, we can show that $P(\theta^*)$ Pareto dominates $P(\theta_2)$:

- The types that pool under both $P(\theta_2)$ and $P(\theta^*)$ are weakly better off under $P(\theta^*)$.

- The types that separate under both $P(\theta_2)$ and $P(\theta^*)$ take strictly more leave under $P(\theta^*)$.

- We can check that the types that separate under $P(\theta_2)$ and pool under $P(\theta^*)$, and who take less leave than the highest leave level for the separating types under $P(\theta^*)$ prefer $P(\theta^*)$ by looking at the indifference condition for type θ^* under $P(\theta^*)$.

[47] The lexicographically maximal Perfect Bayesian Nash equilibrium in this case is equivalent to the lexicographically maximal sequential equilibrium (LMSE) in Mailath et al. (1993) because there are only two periods in the reduced-form problem (Fudenberg and Tirole, 1991).

- We can check that the types that separate under $P(\theta_2)$ and pool under $P(\theta^*)$, and who take more leave than the highest leave level for the separating types under $P(\theta^*)$ prefer $P(\theta^*)$ because, if not, then $P(\theta_2)$ will defeat $P(\theta^*)$, which is a contradiction as the LME is undefeated.

B.3 Proof of Proposition 3

Lemma 1. *The equilibrium in this class in which the most committed type takes the longest possible leave can be obtained by determining the type $\hat{\theta}$ such that $\phi'(\hat{\theta}) = \ell'(\hat{\theta})$, as long as $\tau > (\theta_{\max} - \theta_{\min})\frac{\left(C'\left(-\frac{\tau \log \theta_{\max}}{\theta_{\max}-\theta_{\min}}+K\right)\right)^2}{\min_{\theta \in [\theta_{\min}, \theta_{\max}]} C''\left(-\frac{\tau \log \theta}{\theta_{\max}-\theta_{\min}}+K\right)}$.*

Proof of Lemma. To establish this, it suffices to show that $\phi(\theta)$ is concave and $\ell(\theta)$ is convex.

To show that $\phi(\theta)$ is concave, note that differentiating $\phi(\theta)$ gives

$$\frac{d\phi}{d\theta} = -\left(C^{-1}\right)'(B)\frac{dA}{d\theta}$$

where $B = C(1 - \ell_{\max}) + A$ and $A = \frac{\tau}{\theta}\frac{\theta_{\max}-\theta}{\theta_{\max}-\theta_{\min}}$, and hence

$$\frac{d^2\phi}{d\theta^2} = \left(C^{-1}\right)''(B)\left(\frac{dA}{d\theta}\right)^2 - \left(C^{-1}\right)'(B)\frac{d^2A}{d\theta}.$$

Since $\frac{dA}{d\theta} < 0$ and $\frac{d^2A}{d\theta^2} > 0$, we conclude $\frac{d^2\phi}{d\theta^2} < 0$.

Next, to show that $\ell(\theta)$ is convex, note that

$$
\begin{aligned}
\frac{d\ell}{d\theta} &= \frac{\tau}{\theta(\theta_{\max}-\theta_{\min})}\left(C^{-1}\right)'\left(-\frac{\tau\log\theta}{\theta_{\max}-\theta_{\min}}+K\right) > 0\\
\frac{d^2\ell}{d\theta} &= \frac{-\tau^2}{\theta^2(\theta_{\max}-\theta_{\min})^2}\left(C^{-1}\right)''\left(-\frac{\tau\log\theta}{\theta_{\max}-\theta_{\min}}+K\right)\\
&\quad-\frac{\tau}{\theta^2(\theta_{\max}-\theta_{\min})}\left(C^{-1}\right)'\left(-\frac{\tau\log\theta}{\theta_{\max}-\theta_{\min}}+K\right)\\
&= \frac{-\tau^2}{\theta^2(\theta_{\max}-\theta_{\min})^2}\left(-\frac{C''\left(C^{-1}\left(-\frac{\tau\log\theta}{\theta_{\max}-\theta_{\max}}+K\right)\right)}{\left(C'\left(C^{-1}\left(-\frac{\tau\log\theta}{\theta_{\max}-\theta_{\max}}+K\right)\right)\right)^3}\right)\\
&\quad-\frac{\tau}{\theta^2(\theta_{\max}-\theta_{\min})}\frac{1}{C'\left(C^{-1}\left(-\frac{\tau\log\theta}{\theta_{\max}-\theta_{\max}}+K\right)\right)}\\
&= \frac{\tau}{\theta^2(\theta_{\max}-\theta_{\min})^2}\frac{1}{\left(C'\left(C^{-1}\left(-\frac{\tau\log\theta}{\theta_{\max}-\theta_{\max}}+K\right)\right)\right)}\left[\tau C''\left(C^{-1}\left(\frac{-\tau\log\theta}{\theta_{\max}-\theta_{\min}}+K\right)\right)\right.\\
&\quad\left.-(\theta_{\max}-\theta_{\min})C'\left(C^{-1}\left(\frac{-\tau\log\theta}{\theta_{\max}-\theta_{\min}}+K\right)\right)\right]\\
&> 0
\end{aligned}
$$

as long as τ is sufficiently large so that the term in brackets is positive. □

(1) Fully separating

Since wages depend on beliefs, and since separating equilibria reveal beliefs perfectly, wages do not change if the equilibrium is fully separating both before and after the leave extension.

(2) Partial pooling

To show that the fraction of poolers decreases as $\ell_{\max}$ is extended, it suffices to show that $\frac{d\theta}{d\ell_{\max}} > 0$. Note that $\frac{d\ell}{d\theta} = \frac{d\phi}{d\theta}$ implies

$$\frac{\tau}{\theta(\theta_{\max}-\theta_{\min})}\left(C^{-1}\right)'\left(-\frac{\tau\log\theta}{\theta_{\max}-\theta_{\min}}+K\right) = -\left(C^{-1}\right)'(B)\frac{dA}{d\theta}$$

and differentiating this with respect to $\ell_{\max}$ gives

$$\begin{aligned}\frac{d^2\ell}{d\theta^2}\frac{d\theta}{d\ell^{\max}} &= \left(C^{-1}\right)''(B)\left(-C'(1-\ell_{\max})+\frac{dA}{d\theta}\frac{d\theta}{d\ell_{\max}}\right)\left(-\frac{dA}{d\theta}\right)\\ &\quad+\left(C^{-1}\right)'(B)\left(-\frac{d^2A}{d\theta^2}\frac{d\theta}{d\ell_{\max}}\right)\end{aligned}$$

Thus

$$\frac{d\theta}{d\ell_{\max}} = \frac{D}{\frac{d^2\ell}{d\theta^2}+E+F}$$

where

$$\begin{aligned}D &= \left(C^{-1}\right)''(C(1-\ell^{\max})+A)(-C'(1-\ell^{\max}))\left(-\frac{dA}{d\theta}\right)\\ E &= -\left(C^{-1}\right)''(B)\frac{dA}{d\theta}\left(-\frac{dA}{d\theta}\right)\\ F &= \left(C^{-1}\right)'(B)\left(\frac{d^2A}{d\theta^2}\right).\end{aligned}$$

Since $D > 0$, $E > 0$, and $F > 0$, this establishes the desired result.

(3) and (4) Pre-poolers

The pre-poolers consist of the types $(\theta^*, \theta_{\max})$. Since the fraction of poolers decreases when $\ell_{\max}$ is increased, there exists a θ_1 such that (θ^*, θ_1) separate (PN group) and $(\theta_1, \theta_{\max})$ continue to pool after leave is extended (PP group). Types $\theta \in (\theta^*, \theta_1)$

and $\theta \in (\theta_1, \theta_{\max})$ are believed to be the expected type in $(\theta^*, \theta_{\max})$ before the leave extension since they choose the same action $\ell_{\max}$. After the leave extension, the belief about the former group improves and becomes θ (due to separation). The belief about the latter group becomes worse since $\theta \in (\theta_1, \theta_{\max})$ is pooling with a worse set of types. Wages change accordingly, with increases for the PN group and decreases for the PP group.

C Data appendix

C.1 Leave duration

The leave duration variable comes from two main sources: The SHSS (Sammenhængende socialstatistik) register from 1984 to 2007, and the Ministry of Employment's DREAM database from 1991 to 2012.

The SHSS register contains information on income-based benefits for each month of the year, at the individual level, including temporary benefits such as unemployment benefit, sickness benefit, maternity benefit, cash benefits, and others. The main variables of interest are the VARMMSF variables, recording for each month of the year, the number of days an individual receives maternity allowance. While the number of days is precisely recorded, when an individual is on benefit for a full month, it is recorded as 30 days. I correct for this by adjusting any period of 30 days prior to the last month of a consecutive block of benefits to the correct length of the month. A parental leave duration period is counted starting on the first day a child is born until the parent is off of the benefit for a minimum of 3 weeks. Less than 5% of parents exceed the maximum allowed duration, potentially because of exception rules in the case of unexpected health problems, and they are not included in the analysis. Because the data are only available starting on 1 January 1984, I am only able to use 85 days to the left of the 1984 reform cutoff date of 25 March 1984. The 85-day window is the main specification for the reported results. Appendix Figure 9 shows that the main results are not sensitive to the choice of the window size or the size of the donut hole around the cutoff date.

The DREAM register provides individual-level information on the main benefit type received each week starting in 1991. The benefit types include unemployment benefit, sickness benefit, maternity benefit, child care benefit, cash benefits, education benefits, among others. Even if the benefit is received for one day of the week, it is recorded. However, if multiple benefits are received in the same week, the one with the highest amount is recorded. The data cover the period of child care leave, and also provide additional information on parental leave after 2007. In the analysis in Section 2, child care leave is included if it is used within the first two years when a child is born.

C.2 Labor market data

The IDA data include three sub-panels, IDA-P (personal information), IND-N (employment information), IDA-S (workplace information).

IDA-P includes all individuals aged between 15 and 74 residing in Denmark on the last day of a given year. The variables used include:

- Age: from the Central Danish Person register (CPR)
- Gender: from the Central Danish Person register (CPR)
- Education level: from the administrative educational register (UDDA). The raw education code is at a 6-digit level, but it is aggregated up into three main levels: primary education, high school and vocational education, and further than high school or vocational education. There is a small fraction with unidentified education level, most often for immigrants migrating to Denmark after finishing their education, and this is treated as a separate category.
- County of residence: A unique identifier of the 17 counties.
- Marital status: married, cohabiting, or single.

IDA-N provides labor market information based on annual tax filings from Danish tax authorities. The variables are recorded as of 28 November of a given year. The unit of observation is person-year. The variables include:

- Employment status (`pstill`): Detailed information on the type of employment as of November of the year, including self-employment, employed but on leave, employed, unemployed, in education or out of the labor force, in retirement, in active training.

- Earnings: Sum of net salaries from all information sheets, reported in DKK and CPI adjusted to 2010

- Hours: Hours worked is estimated by Statistics Denmark using the amount of mandatory ATP pension payments of the job. All employers are responsible for reporting the amount to the Danish Tax Authorities.

- Hourly wage: Is measured for the main November job, by dividing total annual earnings for the main job by the total of hours worked that year for the same job.

- Cumulative work experience: The variable is calculated using the cumulative yearly experience variable `erhver`. Each year, work experience is estimated by Statistics Denmark based on ATP payment as a fraction of the year between 0 and 1.

IDA-S provides information on the employer that is associated with the main November job. The 6-digit workplace industry code variable `persbrc` is grouped into the following groups: Agriculture and mining; Manufacturing, energy, and construction; Sales, services, and transportation; Finance and real estate; Education; Health and social work; and Others.

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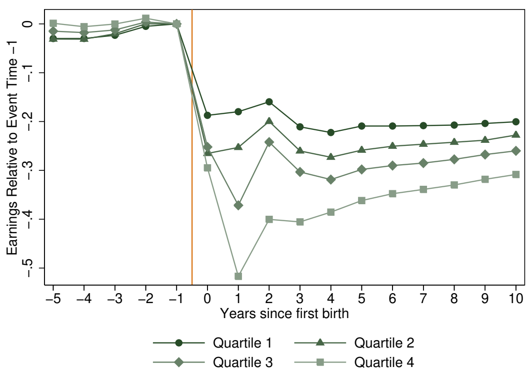
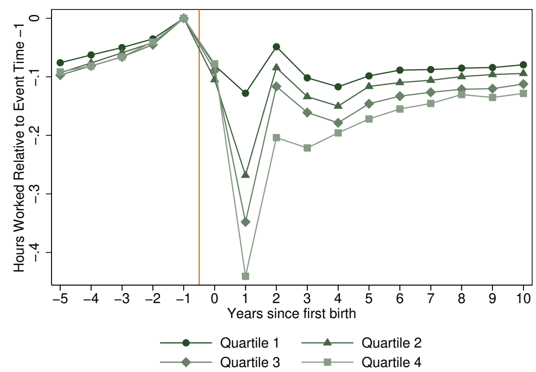
Figure 1: Relative impacts of children by leave-duration quartiles

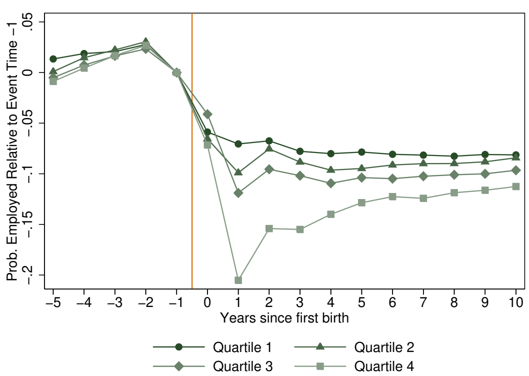
(a) Annual earnings



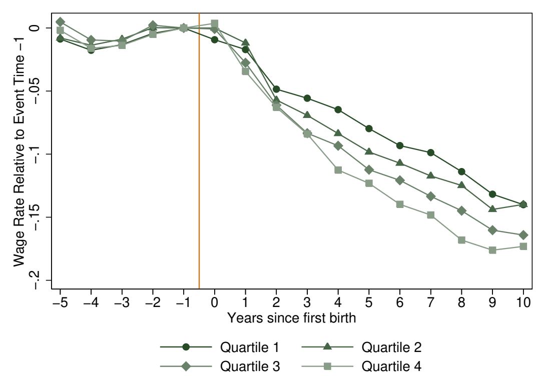
(b) Annual hours worked



(c) Participation rates



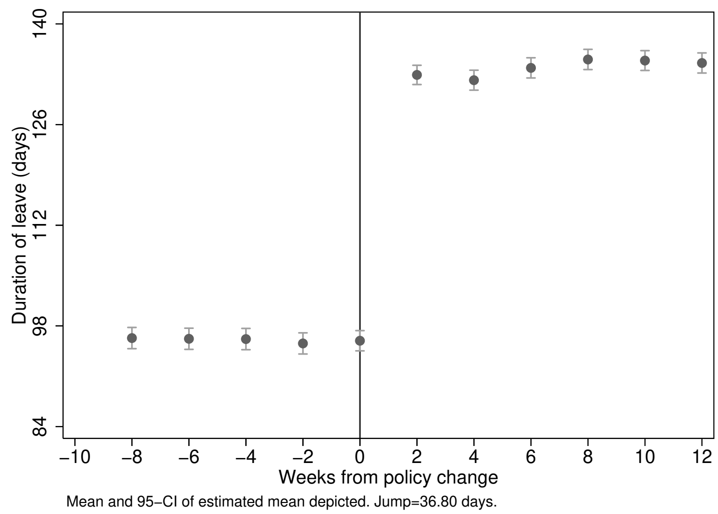
(d) Hourly wage rates



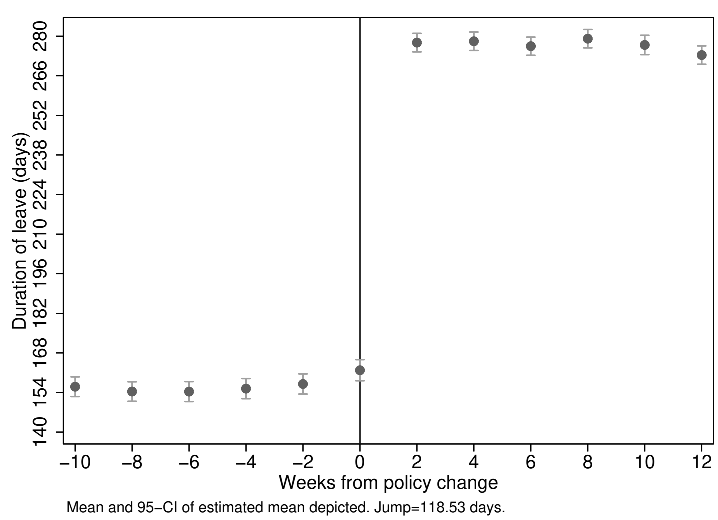
Note: The figure depicts the event time coefficients from equation (1) as a percentage of the counterfactual outcome when children are absent. The coefficients are for mothers relative to fathers, grouped by the cohort-level quartiles (child-birth-year cohort) of leave duration of the mothers. The sample consists of mothers and fathers of 440,605 children born between 1984 and 2003. Annual earning is zero when a parent does not work that year. Annual hours and hourly wage rates are conditional on being employed.

Figure 2: Average leave duration before and after parental leave leave extensions

(a) 25 March 1984 extension



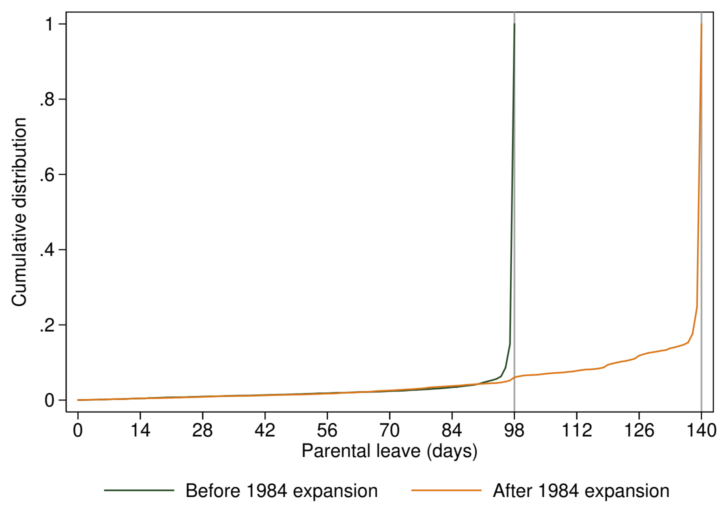
(b) 1 January 2002 extension



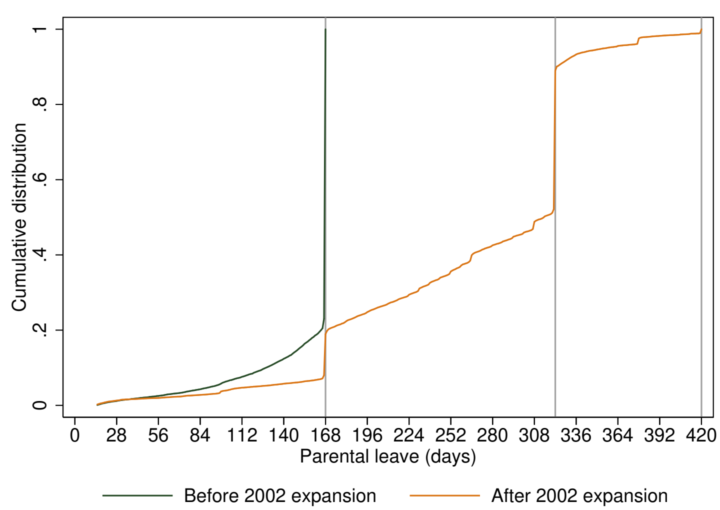
Note: The figure depicts the 95% confidence interval of the mean parental (maternity and shared) leave duration in days, before and after the leave extension in 1984 and 2002. Each point represents a two-week period relative to the policy cutoff dates.

Figure 3: Leave distribution before and after the leave extensions

(a) 25 March 1984 extension



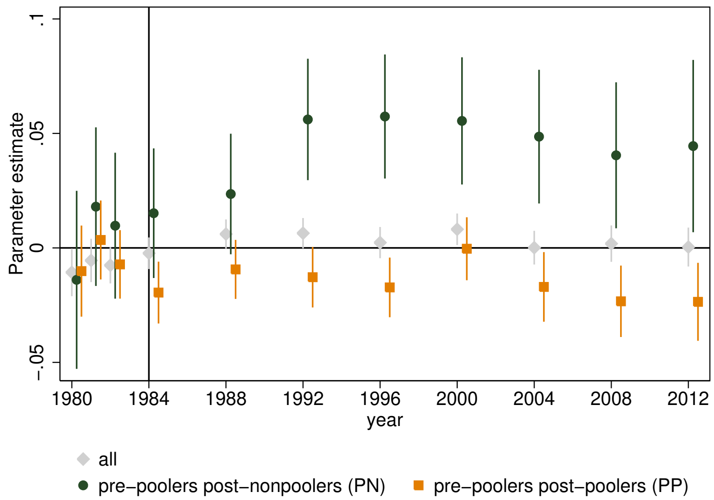
(b) 1 January 2002 extension



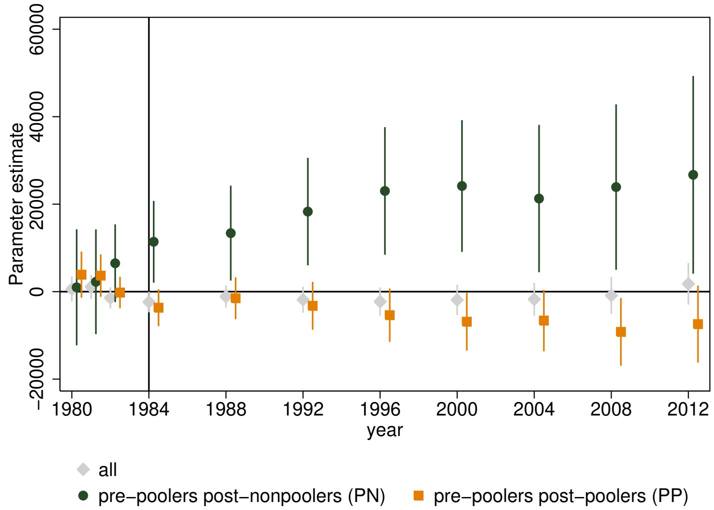
Note: The figure depicts the empirical cumulative distribution of mothers' use of parental (maternity and shared) leave duration in days, before and after the leave extension in 1984 and 2002. In 1984, the window is 85 days on each side of the policy cutoff date. In 2002, the window is 180 days before the policy cutoff date and 180 days after the policy cutoff date.

Figure 4: Dynamic impact of the 1984 leave extension by pooling status

(a) Log wages



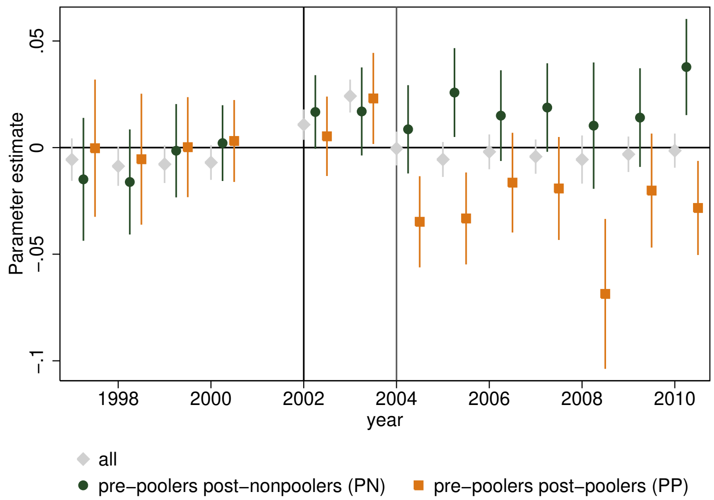
(b) Annual earnings (DKK)



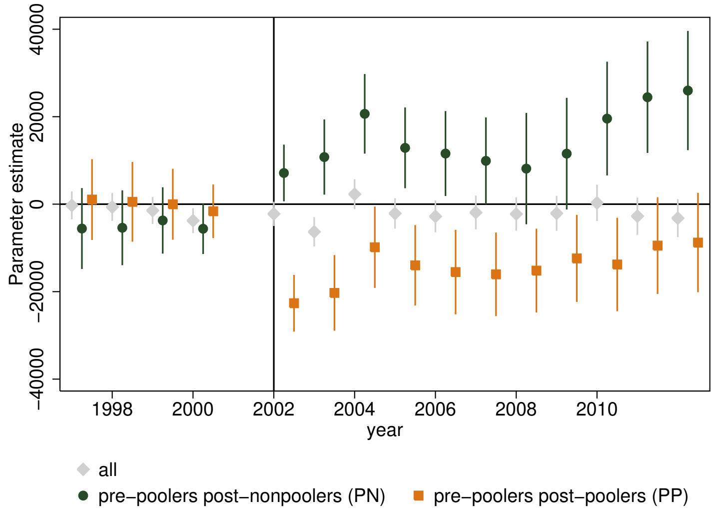
Note: The figure plots the period-by-period birth-after-cutoff coefficients β_k in estimating equation (2) for the 1984 leave extension treatment for three groups of mothers with a child born between 1 January 1984 and 18 June 1984. The three groups are: All mothers with childbirths within this window (All), mothers in the matched pre-pooler post-nonpooler sample (PN), and mothers in the matched pre-pooler post-pooler sample (PP). See Section 4.2 for details. Panel (a) considers log hourly wages as the outcome, and panel (b) considers annual earnings (DKK in 2010) as the outcome. The plotted 95% confidence intervals use standard errors clustered at the individual level.

Figure 5: Dynamic impact of the 2002 extension by pooling status

(a) Log wages

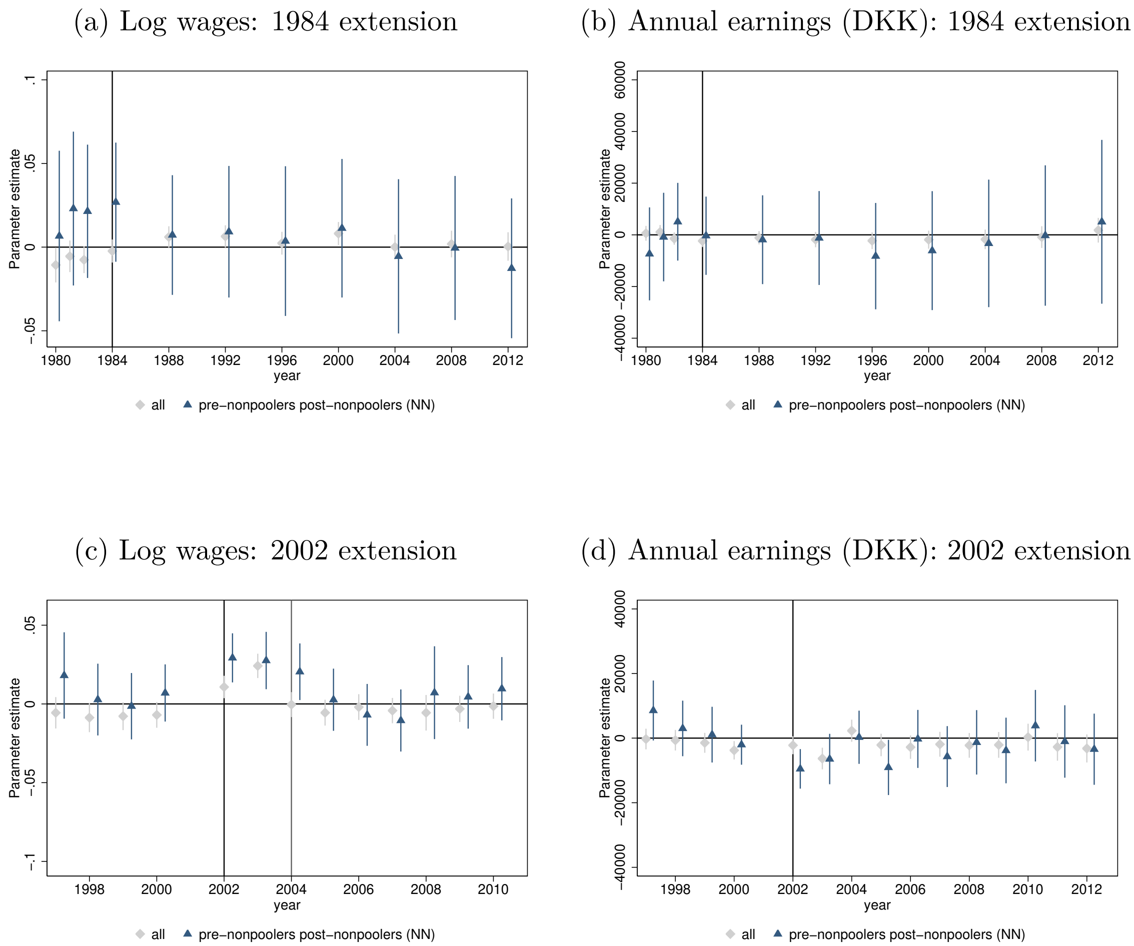


(b) Annual earnings (DKK)



Note: The figure plots the period-by-period birth-after-cutoff coefficients β_k in estimating equation (2) for the 2002 leave extension treatment for three groups of mothers with a child born between 8 October 2001 and 27 March 2002. The three groups are: All mothers with childbirths within this window (All), mothers in the matched pre-pooler post-nonpooler sample (PN), and mothers in the matched pre-pooler post-pooler sample (PP). See Section 4.2 for details. Panel (a) considers log hourly wages as the outcome, and panel (b) considers annual earnings (DKK in 2010) as the outcome. The plotted 95% confidence intervals use standard errors clustered at the individual level.

Figure 6: Dynamic impact of the leave extensions for pre-nonpooler post-nonpoolers



(a) Log wages: 1984 extension

(b) Annual earnings (DKK): 1984 extension

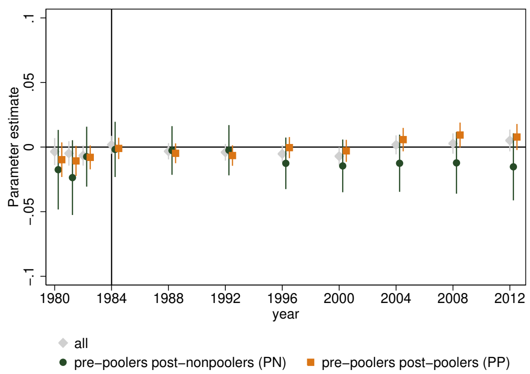
(c) Log wages: 2002 extension

(d) Annual earnings (DKK): 2002 extension

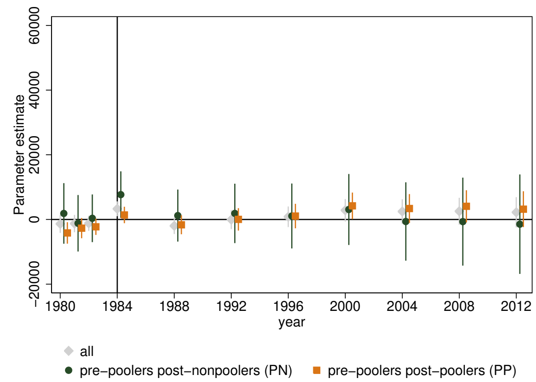
Note: The figure plots the period-by-period birth-after-cutoff coefficients β_k in estimating equation (2) for the two leave extensions treatment for the pre-nonpooler post-nonpooler (NN) mothers. See Section 4.2 for details. Panels (a) and (b) consider the 1984 extension around 25 March 1984 and panels (c) and (d) consider the 2002 extension around 1 January 2002. The plotted 95% confidence intervals use standard errors clustered at the individual level.

Figure 7: Dynamic impact of the placebo policies by pooling status

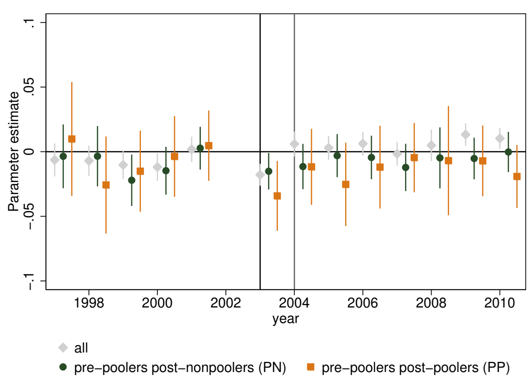
(a) Log wages: 1984 placebo policy



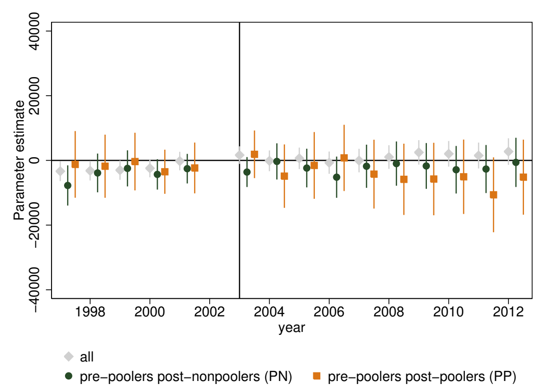
(b) Annual earnings: 1984 placebo policy



(c) Log wages: 2002 placebo policy



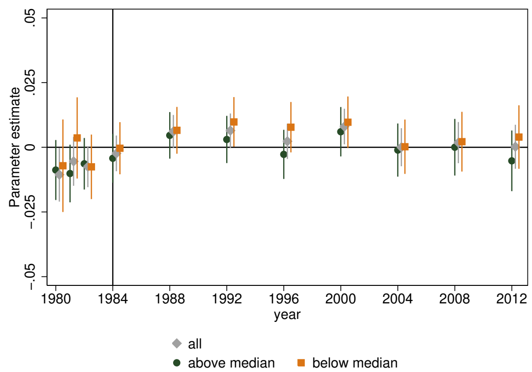
(d) Annual earnings: 2002 placebo policy



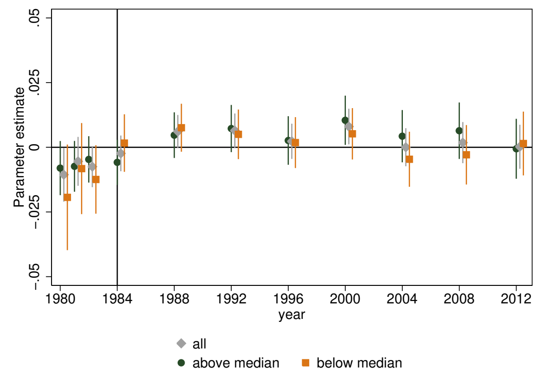
Note: The figure plots the period-by-period birth-after-cutoff coefficients β_k in estimating equation (2) for the 1984 and 2002 placebo policies for three groups of mothers with a child born within a three-month window of the placebo cutoff dates. The 1984 placebo cutoff date is 20 June 1984, and the 2002 placebo cutoff date is 25 June 2002. The three groups are: All mothers with childbirths within this window (All), mothers in the matched pre-pooler post-nonpooler sample (PN), and mothers in the matched pre-pooler post-pooler sample (PP). See Section 4.2 for details. Panels (a) and (b) consider the 1984 placebo policy, and panels (c) and (d) consider the 2002 placebo policy. The plotted 95% confidence intervals use standard errors clustered at the individual level.

Figure 8: Dynamic impact of the 1984 extension by characteristics

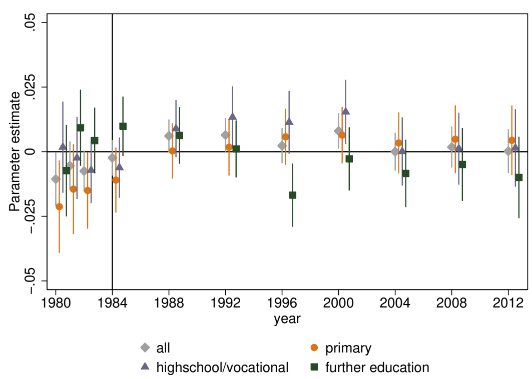
(a) Log wages: By age



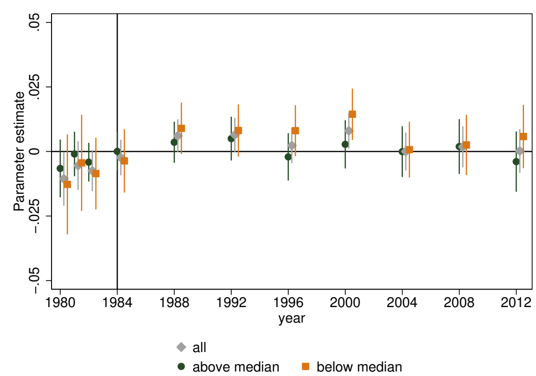
(b) Log wages: By years of experience



(c) Log wages: By education groups



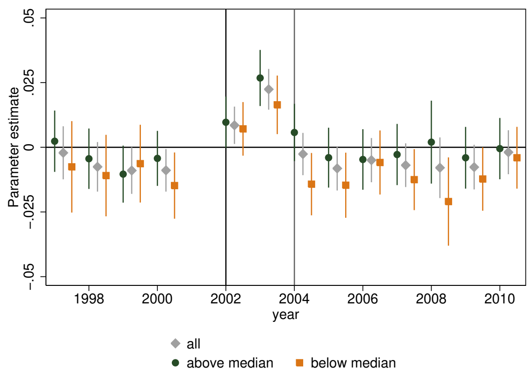
(d) Log wages: By pre-birth annual earnings



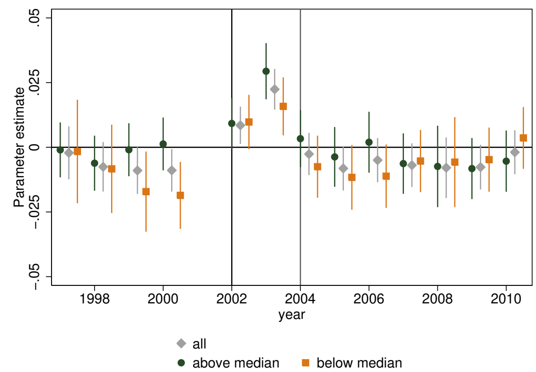
Note: The figure plots the period-by-period birth-after-cutoff coefficients β_k in estimating equation (2) on log hourly wages for the 1984 leave extension treatment by pre-birth characteristics for mothers with a child born between 1 January 1984 and 18 June 1984 and who were working in 1982. See Section 4.2 for details. Panel (a) plots the policy impact for mothers whose age in 1982 were above and below median. Panel (b) plots the policy impact for mothers whose years of experience in 1982 were above and below median. Panel (c) plots the policy impact for mothers whose highest degree is primary school, high school/vocational school, and further education. Panel (d) plots the policy impact for mothers whose annual earnings in 1982 were above and below median. The plotted 95% confidence intervals use standard errors clustered at the individual level.

Figure 9: Dynamic impact of the 2002 extension by characteristics

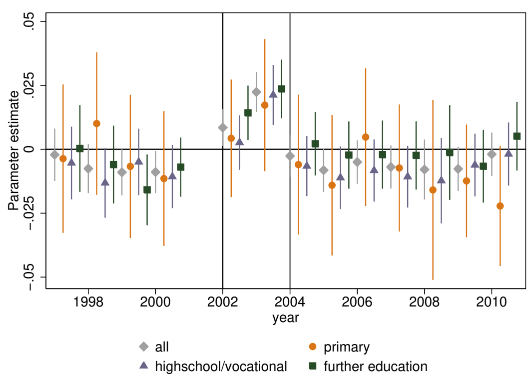
(a) Log wages: By age



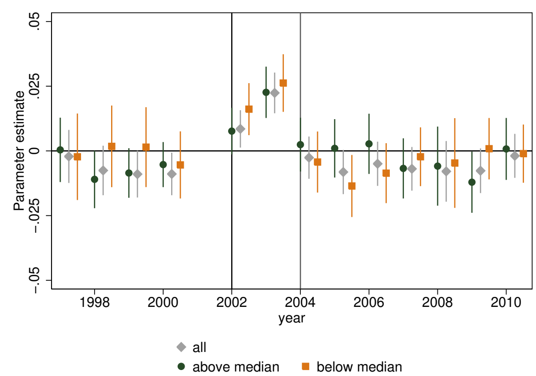
(b) Log wages: By years of experience



(c) Log wages: By education groups



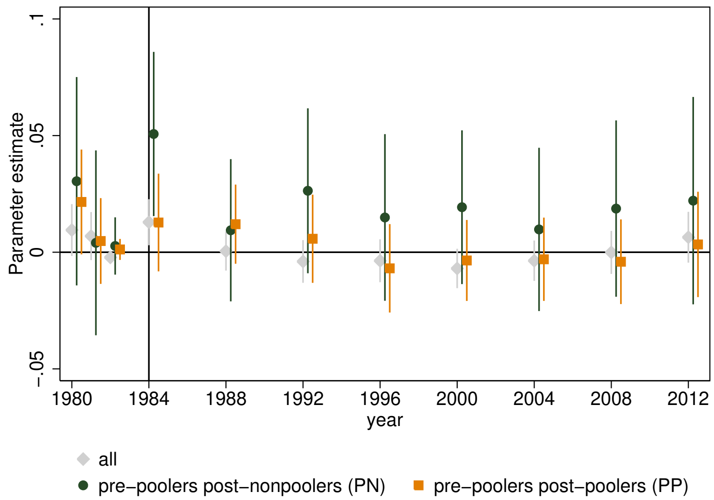
(d) Log wages: By pre-birth annual earnings



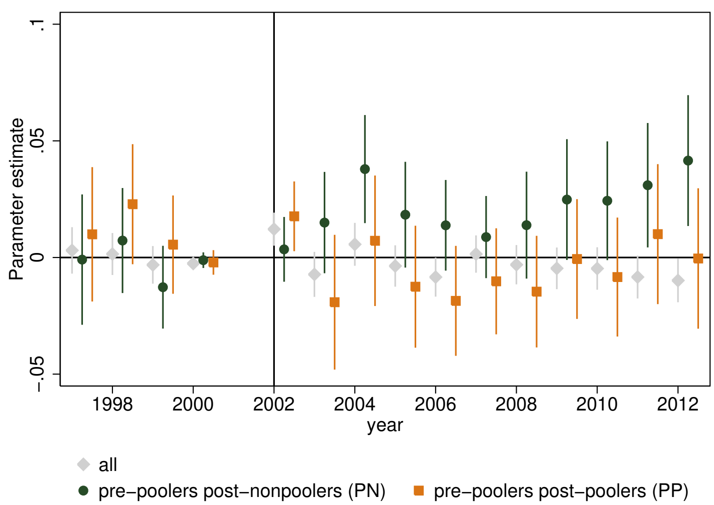
Note: The figure plots the period-by-period birth-after-cutoff coefficients β_k in estimating equation (2) on log hourly wages for the 2002 leave extension treatment by pre-birth characteristics for mothers with a child born between 8 October 2001 and 27 March 2002 and who were working in 1982. See Section 4.2 for details. Panel (a) plots the policy impact for mothers whose age in 1982 were above and below median. Panel (b) plots the policy impact for mothers whose years of experience in 1982 were above and below median. Panel (c) plots the policy impact for mothers who highest degree is primary school, high school/vocational school, and further education. Panel (d) plots the policy impact for mothers whose annual earnings in 1982 were above and below median. The plotted 95% confidence intervals use standard errors clustered at the individual level.

Figure 10: Dynamic impact of the leave extensions on employment status

(a) Employment status: 1984 extension

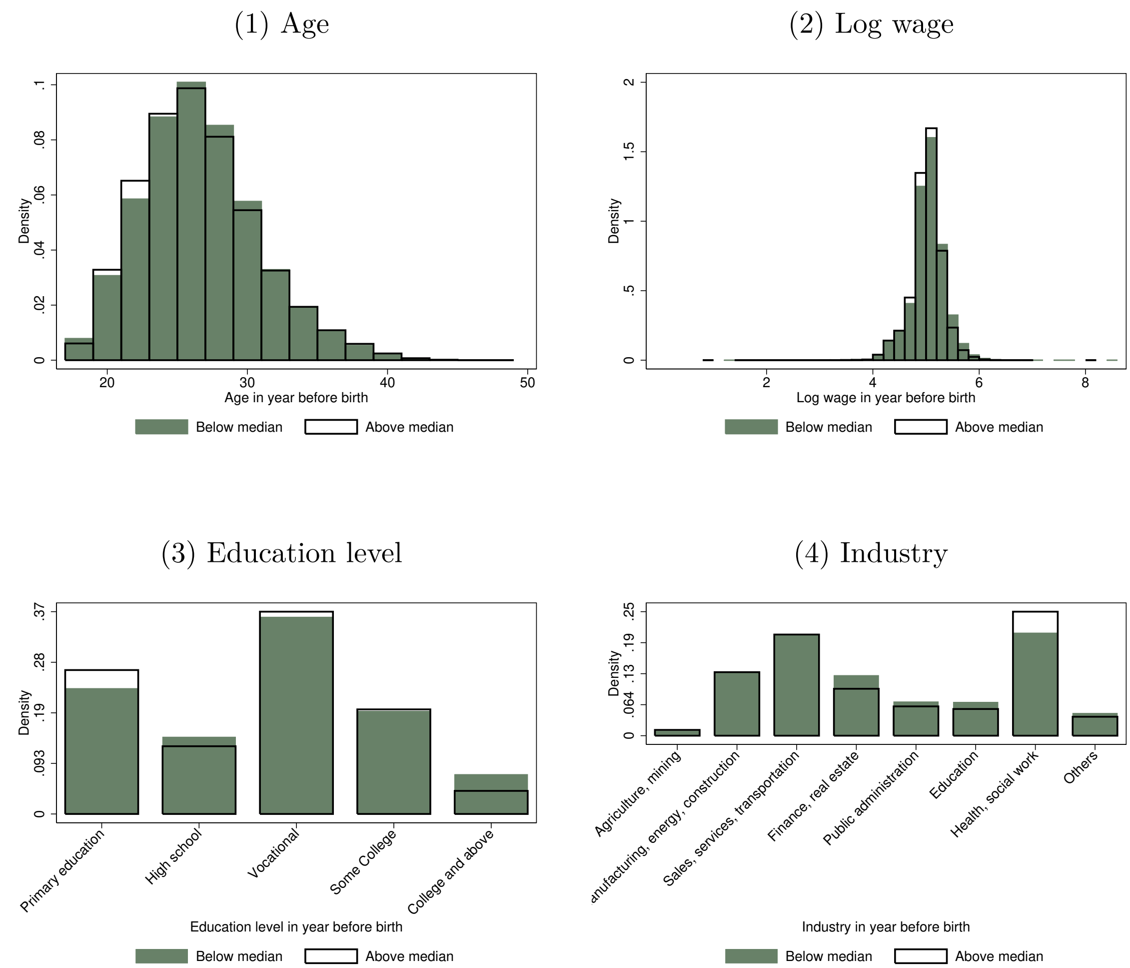


(b) Employment status: 2002 extension



Note: The figure plots the period-by-period birth-after-cutoff coefficients β_k in estimating equation (2) for the 1984 and 2002 leave extension treatment on employment status for three groups of mothers. The three groups are: All mothers with childbirths within this window (All), mothers in the matched pre-pooler post-nonpooler sample (PN), and mothers in the matched pre-pooler post-pooler sample (PP). See Section 4.2 for details. Panel (a) considers the 1984 leave extension around 25 March 1984, and panel (b) considers the 2002 leave extension around 1 January 2002. The plotted 95% confidence intervals use standard errors clustered at the individual level.

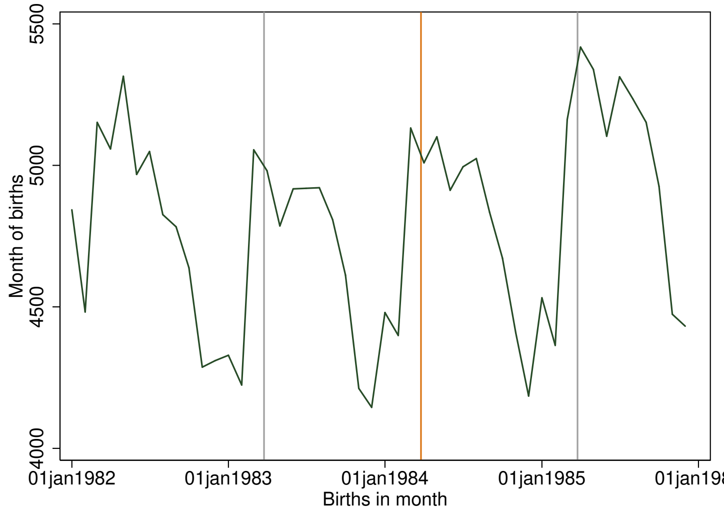
Appendix Figure 1: Distributions of pre-birth characteristics by leave duration



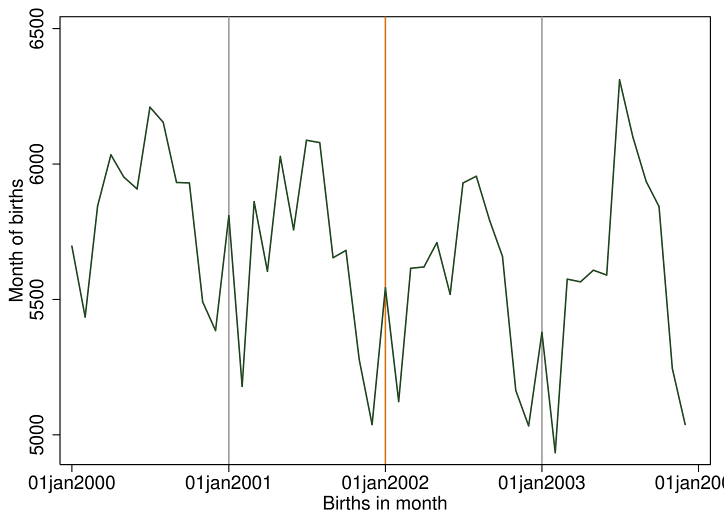
Note: The figure depicts the distributions of pre-birth characteristics of first-time mothers whose leave durations are below or above the median compared to mothers giving birth in the same year. The sample consists of mothers and fathers of 440,605 children born between 1984 and 2003. The solid green blocks represent the distributions when leave duration is below the median. The hollow blocks represent the distributions when leave duration is above the median.

Appendix Figure 2: Birth seasonality around 1984 and 2002

(1) Monthly birth counts 1982—1985



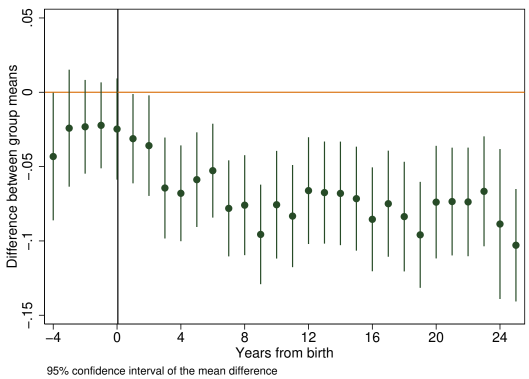
(2) Monthly birth counts 2000—2003



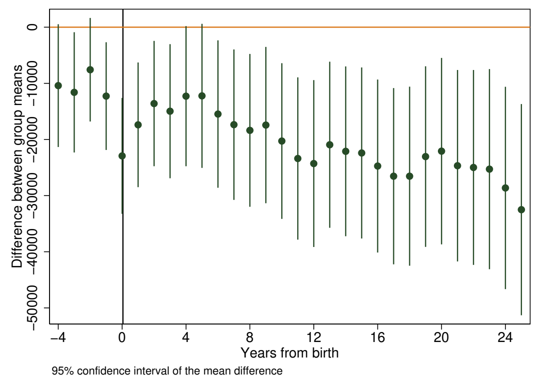
Note: The figure depicts total number of births each month in the periods 1982 to 1985 and in the periods 2000 to 2003.

Appendix Figure 3: Raw mean differences between poolers and nonpoolers after 1984 extension

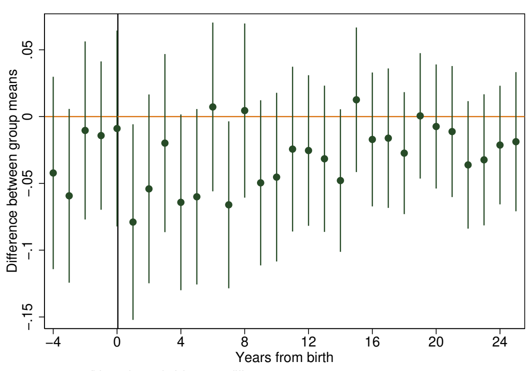
(3) Log wages



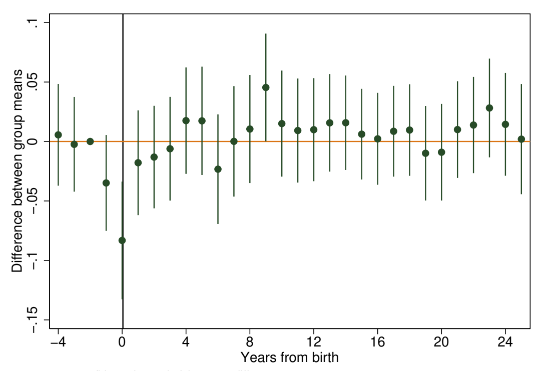
(4) Annual earnings (DKK in 2010)



(5) Log annual hours

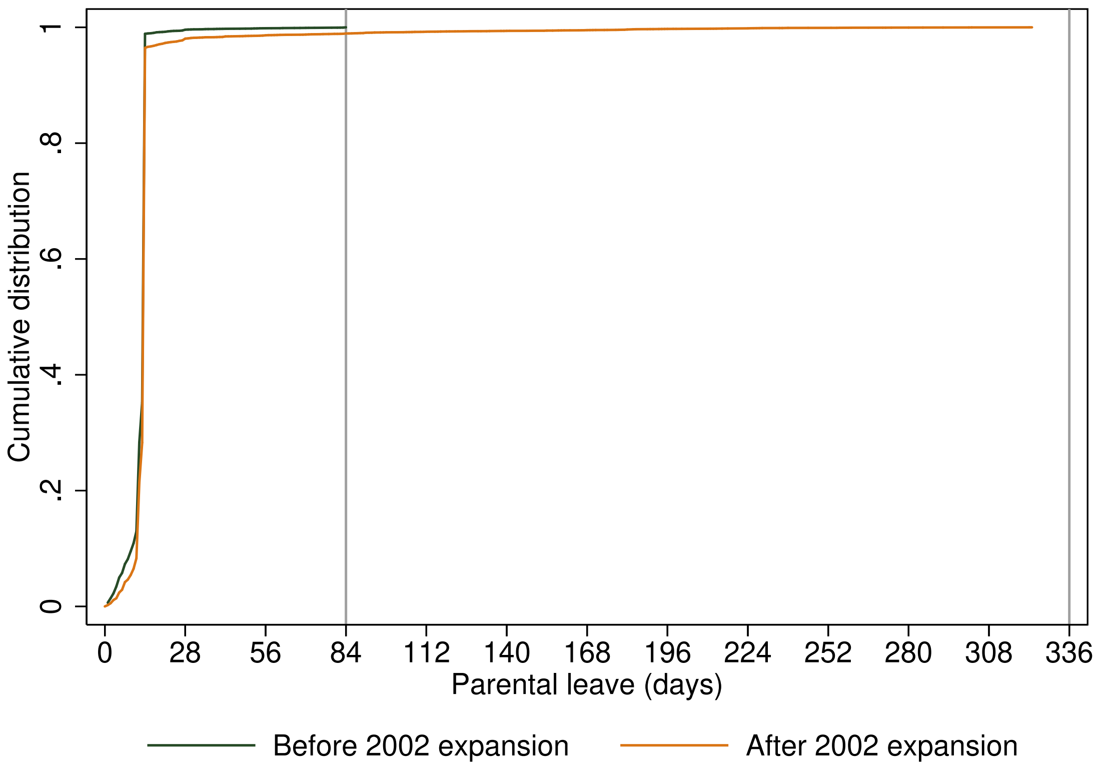


(6) Employment status



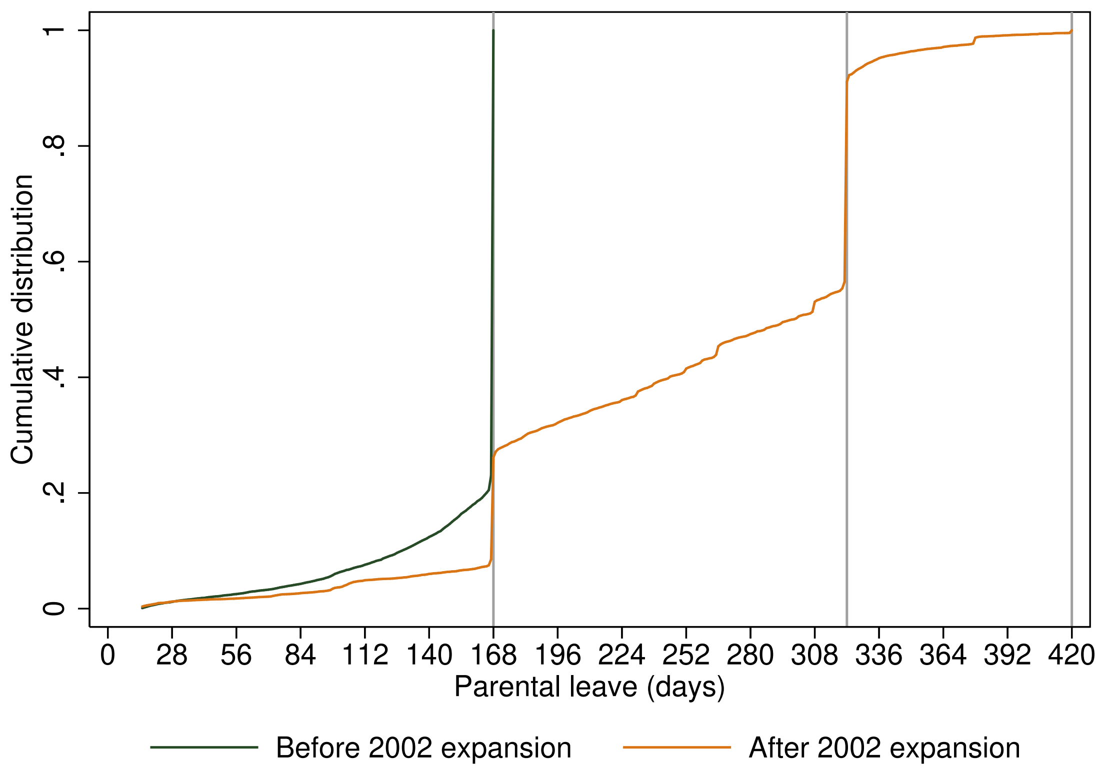
Note: The figure plots the labor market outcome mean differences between mothers who take 140 days of parental leave and mothers who take between 130 and 139 days of parental leave and who had children born between 25 March 1984 and 18 June 1984.

Appendix Figure 4: Leave distribution of fathers before and after the 2002 extension



Note: The figure depicts the empirical cumulative distribution of fathers' use of parental (paternity and shared) leave duration in days, before and after the leave extension in 2002. The window is 180 days before the policy cutoff date and 85 days after the policy cutoff date, the period during which mothers can choose between the two policy options.

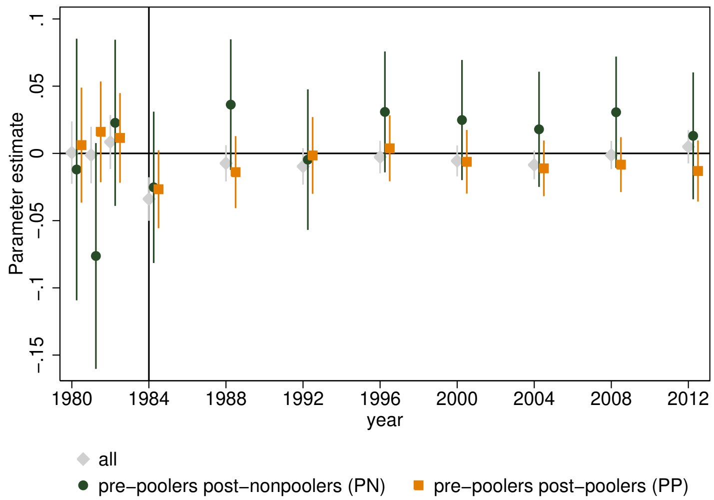
Appendix Figure 5: Leave distribution before and after the 2002 extension



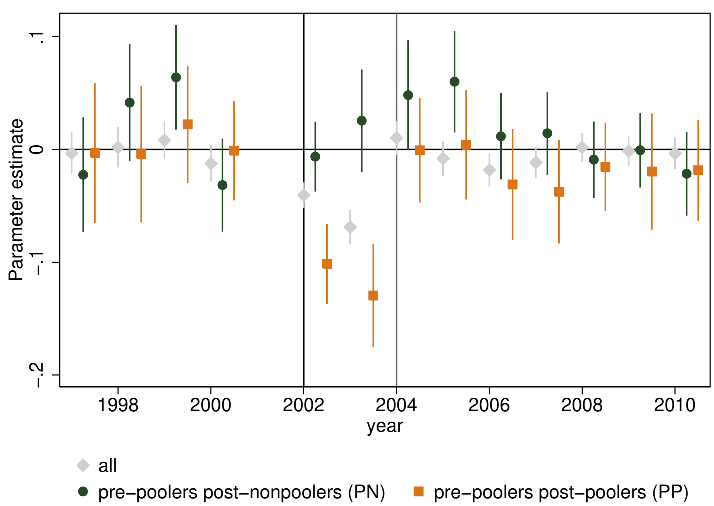
Note: The figure depicts the empirical cumulative distribution of mothers' use of parental (maternity and shared) leave duration in days, before and after the leave extension in 2002. The window is 180 days before the policy cutoff date and 85 days after the policy cutoff date, the period during which mothers can choose between the two policy options.

Appendix Figure 6: Dynamic impact of the leave extensions on log hours

(1) Log hours: 1984 extension



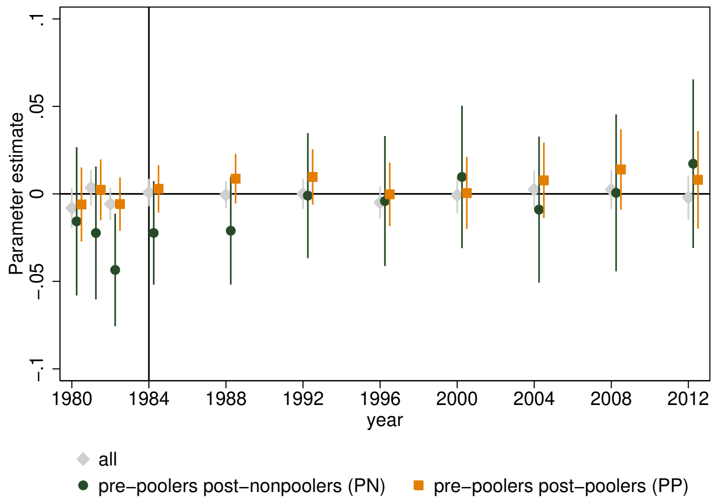
(2) Log hours: 2002 extension



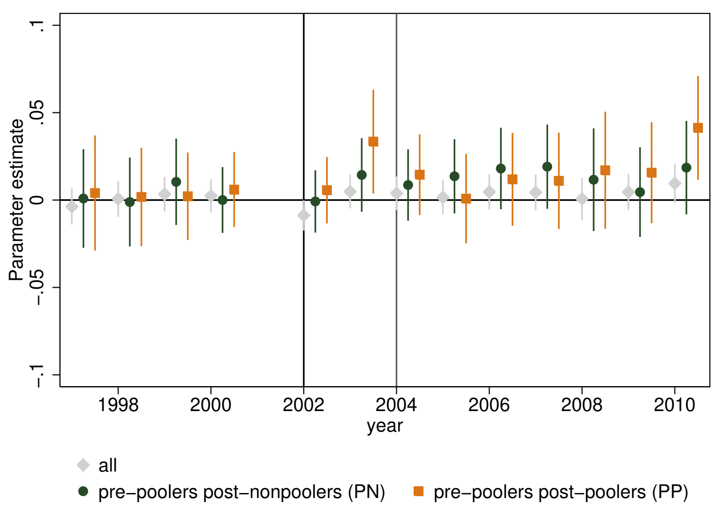
Note: The figure plots the period-by-period birth-after-cutoff coefficients β_k in estimating equation (2) for the 1984 and 2002 leave extension treatment on log annual hours worked for three groups of mothers. The three groups are: All mothers with childbirths within this window (All), mothers in the matched pre-pooler post-nonpooler sample (PN), and mothers in the matched pre-pooler post-pooler sample (PP). See Section 4.2 for details. Panel (a) considers the 1984 leave extension around 25 March 1984, and panel (b) considers the 2002 leave extension around 1 January 2002. The plotted 95% confidence intervals use standard errors clustered at the individual level.

Appendix Figure 7: Dynamic impact of the leave extensions on fathers' wages

(1) Fathers' log wages: 1984 extension



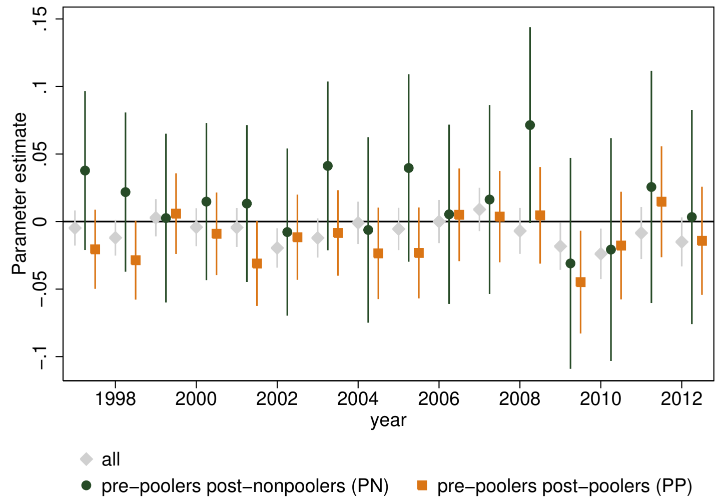
(2) Fathers' log wages: 2002 extension



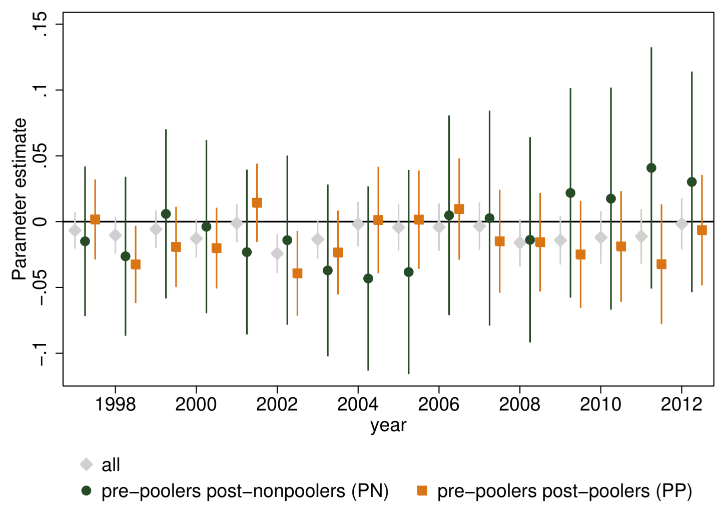
Note: The figure plots the period-by-period birth-after-cutoff coefficients β_k in estimating equation (2) for the 1984 and 2002 leave extension treatment on log hourly wages of the fathers whose partners are in three groups. The three groups are: All mothers with childbirths within this window (All), mothers in the matched pre-pooler post-nonpooler sample (PN), and mothers in the matched pre-pooler post-pooler sample (PP). See Section 4.2 for details. Panel (a) considers the 1984 leave extension around 25 March 1984, and panel (b) considers the 2002 leave extension around 1 January 2002. The plotted 95% confidence intervals use standard errors clustered at the individual level.

Appendix Figure 8: Dynamic impact of the 1984 extension on hospital record count

(1) Log number of hospital records (child)

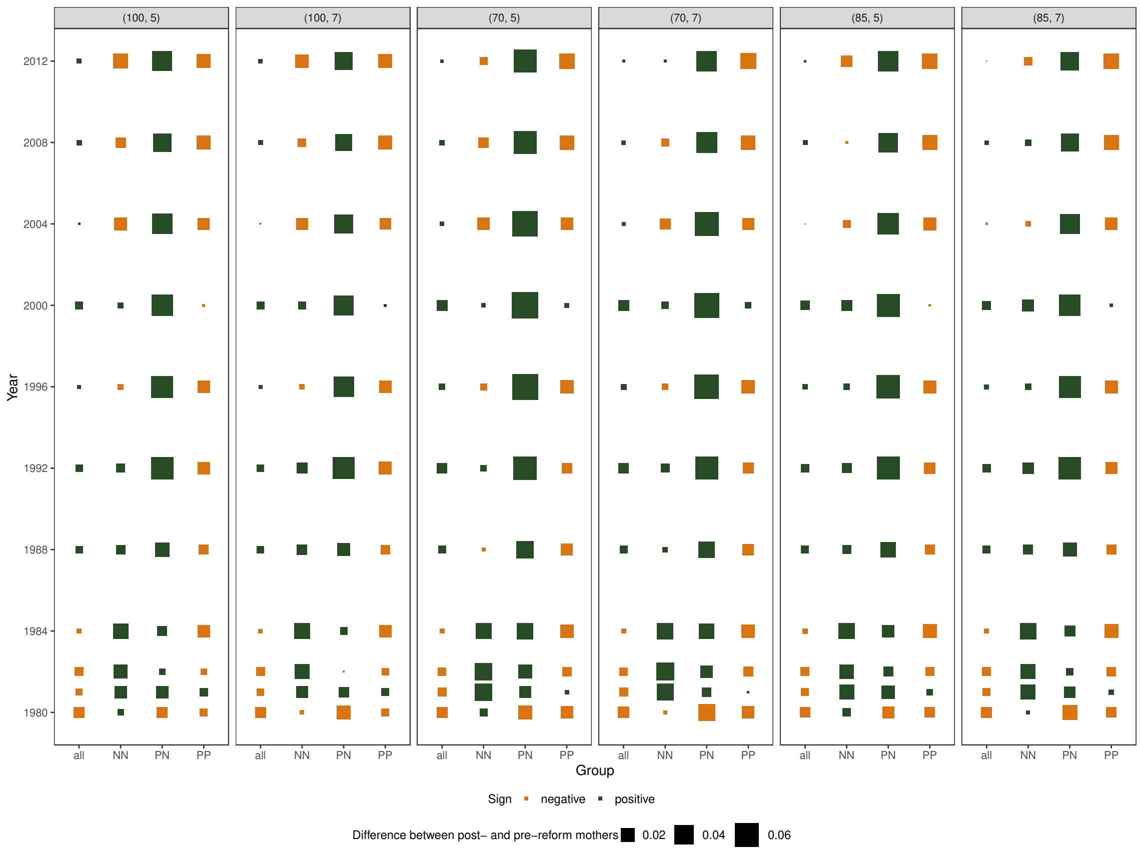


(2) Log number of hospital records (mother)



Note: The figure plots the period-by-period birth-after-cutoff coefficients β_k in estimating equation (2) for the 1984 leave extension treatment for three groups of mothers with a child born between 1 January 1984 and 18 June 1984. The three groups are: All mothers with childbirths within this window (All), mothers in the matched pre-pooler post-nonpooler sample (PN), and mothers in the matched pre-pooler post-pooler sample (PP). See Section 4.2 for details. Panel (a) considers log number of hospital records for the child as the outcome, and panel (b) considers log number of hospital records for the mother as the outcome. Hospital record count is available starting in 1997. The plotted 95% confidence intervals use standard errors clustered at the individual level.

Appendix Figure 9: Dynamic impact of the 1984 leave extension on log wages by pooling status: Varying window sizes



Note: The figure plots the period-by-period birth-after-cutoff coefficients β_k in estimating equation (2) for the 1984 leave extension treatment for four groups of mothers with a child born around 25 March 1984. The four groups are: All mothers with childbirths within this window (All), mothers in the matched pre-nonpooler post-nonpooler sample (NN), mothers in the matched pre-pooler post-nonpooler sample (PN), and mothers in the matched pre-pooler post-pooler sample (PP). See Section 4.2 for details. The outcome is log hourly wages. Each panel (w, d) represents the specification using a window size of w days on each side of the cutoff date (up to 85 days on the left side of the cutoff date due to data availability), and with a donut hole of d days around the cutoff date.

Table 1: Pre-birth group characteristics around the 1984 extension

Characteristic	All	PN	PP	NN
Mother's annual earnings (DKK)	196,334.3	225,261.4	203,140.2	237,278.8
	(2152.0)	(893.4)	(898.9)	(3772.7)
Father's annual earnings (DKK)	246,537.1	272,510.6	259,490.3	283,741.0
	(1081.7)	(3732.9)	(1846.9)	(5790.8)
Age of mother	26.16	26.54	26.06	27.43
	(0.034)	(0.115)	(0.0567)	(0.169)
Age of father	28.55	28.54	28.30	29.37
	(0.047)	(0.120)	(0.0628)	(0.175)
Mother with high school/vocational (%)	0.346	0.320	0.395	0.246
	(0.0036)	(0.0134)	(0.0072)	(0.0164)
Father with high school/vocational (%)	0.451	0.459	0.611	0.362
	(0.0037)	(0.0143)	(0.0072)	(0.0183)
Mother with further education (%)	0.256	0.415	0.213	0.522
	(0.0033)	(0.0141)	(0.0060)	(0.0190)
Father with further education (%)	0.200	0.340	0.158	0.451
	(0.003)	(0.0136)	(0.00535)	(0.0189)
Mother's work experience (years)	5.394	5.646	5.860	5.688
	(0.0259)	(0.0852)	(0.0458)	(0.132)
Father's work experience (years)	6.916	6.698	7.141	6.553
	(0.0293)	(0.0898)	(0.0490)	(0.128)
Mother in public sector (%)	0.509	0.603	0.507	0.602
	(0.0038)	(0.0142)	(0.0074)	(0.019)
Mother working full-time (%)	0.420	0.533	0.449	0.548
	(0.0038)	(0.0144)	(0.0074)	(0.019)
Mother's transfer income (DKK)	2,909.3	1,478.3	1,934.4	1853.3
	(48.67)	(126.2)	(72.52)	(199.8)
Number of observations	17749	1216	4654	692

Note: The table presents summary statistics of the characteristics in 1982 for four groups of mothers, all with a child born between 1 January 1984 and 18 June 1984. The four groups are: All mothers with childbirths within this window (All), mothers in the matched pre-pooler post-nonpooler sample (PN), mothers in the matched pre-pooler post-pooler sample (PP), and mothers in the matched pre-nonpooler post-nonpooler sample (NN). See Section 4.2 for details. Annual earnings and transfer income are measured in DKK in 2010. Further education is defined as having at least two years of schooling after high school/vocational level. Transfer income is the total amount of government provided benefits including cash benefits, unemployment benefits, holiday pay, housing support, and child allowance. Standard errors are in parentheses.

Table 2: Pre-birth group characteristics around the 2002 extension

Characteristic	All	PN	PP	NN
Mother's annual earnings (DKK)	223,870.1	244,713.4	228,783.9	261,637.1
	(821.9)	(1502.6)	(1506.7)	(1568.3)
Father's annual earnings (DKK)	312,024.2	323,819.6	312,298.1	336,453.7
	(1374.0)	(2953.1)	(3184.6)	(2898.5)
Age of mother	29.08	28.36	27.98	29.02
	(0.031)	(0.063)	(0.0716)	(0.0585)
Age of father	30.50	29.90	29.68	30.51
	(0.051)	(0.083)	(0.106)	(0.0776)
Mother with high school/vocational (%)	0.468	0.514	0.578	0.511
	(0.0034)	(0.0090)	(0.0103)	(0.0081)
Father with high school/vocational (%)	0.444	0.552	0.638	0.377
	(0.0034)	(0.0090)	(0.0100)	(0.00786)
Mother with further education (%)	0.359	0.418	0.330	0.484
	(0.0033)	(0.0089)	(0.0098)	(0.0081)
Father with further education (%)	0.286	0.325	0.238	0.377
	(0.003)	(0.0085)	(0.0089)	(0.00786)
Mother's years of experience	6.566	6.571	6.585	6.774
	(0.0287)	(0.0614)	(0.0680)	(0.0587)
Father's years of experience	9.411	9.095	9.392	9.195
	(0.0366)	(0.0778)	(0.0892)	(0.0731)
Mother in public sector (%)	0.420	0.463	0.483	0.431
	(0.0027)	(0.0094)	(0.0109)	(0.0083)
Mother working full-time (%)	0.496	0.567	0.535	0.577
	(0.0036)	(0.0094)	(0.011)	(0.0083)
Mother's transfer income (DKK)	16,448.6	9,825.7	10,150.0	9,544.7
	(179.9)	(286.8)	(368.5)	(265.6)
Number of observations	22035	3056	2310	3805

Note: The table presents summary statistics of the characteristics in 2000 for four groups of mothers, all with a child born between 8 October 2001 and 27 March 2002. The four groups are: All mothers with childbirths within this window (All), mothers in the matched pre-pooler post-nonpooler sample (PN), mothers in the matched pre-pooler post-pooler sample (PP), and mothers in the matched pre-nonpooler post-nonpooler sample (NN). See Section 4.2 for details. Annual earnings and transfer income are measured in DKK in 2010. Further education is defined as having at least two years of schooling after high school/vocational level. Transfer income is the total amount of government provided benefits including cash benefits, unemployment benefits, holiday pay, housing support, and child allowance. Standard errors are in parentheses.

Table 3: Balance of covariates within subsamples for the 1984 extension

Characteristic in 1982	Pre - Post		
	PN	PP	NN
Mother's annual earnings (DKK)	6,065.9	-151.7	-2,036.4
	(0.246)	(0.937)	(0.572)
Father's annual earnings (DKK)	-5,758.0	-665.0	7,559.9
	(0.510)	(0.871)	(0.280)
Age of mother	0.150	0.0249	-0.0318
	(0.580)	(0.840)	(0.815)
Age of father	-0.052	-0.0449	0.0930
	(0.853)	(0.740)	(0.588)
Mother with high school/vocational (%)	-0.0163	0.007	0.0006
	(0.598)	(0.666)	(0.977)
Father with high school/vocational (%)	-0.0125	-0.004	0.0154
	(0.707)	(0.795)	(0.444)
Mother with further education (%)	0.0265	-0.0024	0.0188
	(0.427)	(0.855)	(0.652)
Father with further education (%)	0.0133	-0.0007	0.0139
	(0.681)	(0.952)	(0.740)
Mother's years of experience	-0.0713	0.185	-0.082
	(0.722)	(0.070)	(0.769)
Father's years of experience	-0.110	0.0182	0.254
	(0.599)	(0.868)	(0.342)
Mother in public sector (%)	-0.0231	-0.0167	0.0258
	(0.486)	(0.316)	(0.532)
Mother working full-time (%)	0.0411	0.0082	-0.0219
	(0.224)	(0.621)	(0.601)
Mother's transfer income (DKK)	-37.96	-0.679	-266.5
	(0.898)	(0.996)	(0.580)
Number of observations	1216	4654	692

Note: The table presents tests of balance of the characteristics in 1982 for for three groups of mothers, all with a child born between 1 January 1984 and 18 June 1984. The three groups are: Mothers in the matched pre-pooler post-nonpooler sample (PN), mothers in the matched pre-pooler post-pooler sample (PP), and mothers in the matched pre-nonpooler post-nonpooler sample (NN). See Section 4.2 for details. Annual earnings and transfer income are measured in DKK in 2010. Further education is defined as having at least two years of schooling after high school/vocational level. Transfer income is the total amount of government provided benefits including cash benefits, unemployment benefits, holiday pay, housing support, and child allowance. Differences between the before and after groups are presented, together with p-values in parentheses.

Table 4: Balance of covariates within subsamples for the 2002 extension

Characteristic in 2000	Pre - Post		
	PN	PP	NN
Mother's annual earnings (DKK)	-365.7	1,252.2	-2,036.4
	(0.916)	(0.711)	(0.572)
Father's annual earnings (DKK)	8,739.8	1,257.5	7,559.9
	(0.228)	(0.855)	(0.280)
Age of mother	0.0988	0.0418	-0.0318
	(0.489)	(0.793)	(0.815)
Age of father	0.268	0.0421	0.0930
	(0.151)	(0.851)	(0.588)
Mother with high school/vocational (%)	-0.0047	-0.0174	0.0006
	(0.826)	(0.459)	(0.977)
Father with high school/vocational (%)	-0.0110	0.00346	0.0154
	(0.602)	(0.879)	(0.444)
Mother with further education (%)	0.0106	-0.0185	-0.0057
	(0.615)	(0.413)	(0.778)
Father with further education (%)	0.0250	-0.0063	-0.0103
	(0.217)	(0.758)	(0.594)
Mother's years of experience	-0.0682	0.0713	-0.0060
	(0.629)	(0.645)	(0.966)
Father's years of experience	-0.0887	0.0182	0.185
	(0.658)	(0.868)	(0.304)
Mother in public sector (%)	0.0149	-0.0167	-0.0443
	(0.553)	(0.316)	(0.030)
Mother working full-time(%)	-0.0413	-0.0348	-0.0315
	(0.061)	(0.164)	(0.119)
Mother's transfer income (DKK)	41.81	-1331.0	-197.1
	(0.945)	(0.103)	(0.755)
Number of observations	3056	2310	3805

Note: The table presents tests of balance of the characteristics in 2000 for three groups of mothers, all with a child born between 8 October 2001 and 27 March 2002. The three groups are: Mothers in the matched pre-pooler post-nonpooler sample (PN), mothers in the matched pre-pooler post-pooler sample (PP), and mothers in the matched pre-nonpooler post-nonpooler sample (NN). See Section 4.2 for details. Annual earnings and transfer income are measured in DKK in 2010. Further education is defined as having at least two years of schooling after high school/vocational level. Transfer income is the total amount of government provided benefits including cash benefits, unemployment benefits, holiday pay, housing support, and child allowance. Differences between the before and after groups are presented, together with p-values in parentheses.

Table 5: Wage comparison before/after extensions conditional on leave length

1984			2002		
Year	(1)	(2)	Year	(3)	(4)
1980	0.115	0.0757	1997	0.0379	0.0453
	(0.0174)	(0.0196)		(0.0168)	(0.0178)
1981	0.0941	0.0497	1998	0.0512	0.0587
	(0.0160)	(0.0188)		(0.0166)	(0.0176)
1982	0.0891	0.0485	1999	0.0370	0.0462
	(0.0151)	(0.0180)		(0.0148)	(0.0158)
1984	0.0686	0.0277	2000	0.0319	0.0401
	(0.0127)	(0.0161)		(0.0141)	(0.0152)
1988	0.0965	0.0552	2002	0.0410	0.0506
	(0.0108)	(0.0146)		(0.0132)	(0.0142)
1992	0.113	0.0700	2003	0.0880	0.0972
	(0.0127)	(0.0157)		(0.0135)	(0.0148)
1996	0.147	0.102	2004	0.0530	0.0630
	(0.0136)	(0.0161)		(0.0140)	(0.0154)
2000	0.156	0.112	2005	0.0422	0.0518
	(0.0130)	(0.0159)		(0.0145)	(0.0156)
2004	0.153	0.111	2006	0.0410	0.0502
	(0.0135)	(0.0163)		(0.0148)	(0.0157)
2008	0.140	0.100	2007	0.0512	0.0595
	(0.0143)	(0.0170)		(0.0139)	(0.0152)
2012	0.137	0.0965	2008	0.0527	0.0615
	(0.0152)	(0.0180)		(0.0184)	(0.0197)
Leave length FE	No	Yes	Leave length FE	No	Yes
N	7779		N	7581	

Note: The table presents the differences in log wages of mothers with childbirths around the policy cutoff dates conditional on taking at most the maximum allowed leave duration before the extension took place. For the 1984 policy (columns 1 and 2), the sample includes mothers whose were working in 1982 and who had a child born between 1 January 1984 and 18 June 1984, taking at most 98 days of parental leave. For the 2002 policy (columns 3 and 4), the sample includes mothers whose were working in 2002 and who had a child born between 8 October 2001 and 27 March 2002, taking at most 168 days of parental leave and return to work without using child care leave for at least three weeks. Columns (2) and (4) include fixed effects for leave length measured at the day level. Standard errors are in parentheses and clustered at the individual level.

Table 6: Impact of leave extensions on fertility measures

Outcome	All	PN	PP	NN
Panel A: Impact of 1984 policy change				
Total number of children	0.012	-0.04	0.005	-0.013
	(0.013)	(0.049)	(0.02)	(0.064)
Birth spacing to next child (months)	0.037	-3.01	-0.36	-2.79
	(0.67)	(2.74)	(1.37)	(3.41)
Leave used for next child (days)	0.83	-0.35	0.47	-3.44
	(0.62)	(3.88)	(1.22)	(4.43)
Number of observations	17749	1216	4654	692
Panel B: Impact of 2002 policy change				
Total number of children	0.001	-0.007	0.035	0.029
	(0.012)	(0.028)	(0.034)	(0.026)
Birth spacing to next child (months)	-0.156	-0.658	-0.36	-0.14
	(0.288)	(0.767)	(1.36)	(0.719)
Leave used for next child (days)	2.67	-1.02	-0.43	6.84
	(1.85)	(5.22)	(6.46)	(5.89)
Number of observations	22035	3056	2310	3805

Note: The table presents the birth-after-cutoff coefficients β_1 in estimating equation (3) for the 1984 and 2002 leave extension treatment for four groups of mothers on fertility measures. The four groups are: All mothers with childbirths within this window (All), mothers in the matched pre-pooler post-nonpooler sample (PN), mothers in the matched pre-pooler post-pooler sample (PP), and mothers in the matched pre-nonpooler post-nonpooler sample (NN). See Section 4.2 for details. Panel A considers the 1984 leave extension around 25 March 1984, and panel B considers the 2002 leave extension around 1 January 2002. The three outcomes are: total number of children by 2012, months to the next child if there is any, and parental leave length used for the next child if there is any. Standard errors are clustered at the individual level.

Appendix Table 1: Summary of the 1984 and 2002 leave extensions

1984 extension		2002 extension
October 1984	First proposed	November 2001
December 1984	Passed	(22 March 2002)
25 March 1984	Effective cutoff	1 January 2002
	Changes	
14 to 20 weeks (first 14 not shareable)	Total parental leave extension	24 to 46 weeks* (first 14 not shareable)
Full benefit	Wage replacement	Full benefit
4 more shareable weeks from 11 February 1985	Other change	Remove child care leave (52 weeks at 60% benefit)

*Extendable to 60 weeks with same total pay

Note: The table summarizes the changes impacted by the 1984 parental leave extension and the 2002 parental leave extension. More details can be found in Section 4.1. Fathers are entitled to two weeks of paid leave reserved only for them during the first 14 weeks after a child is born. Mothers giving birth between 1 January 2002 and 26 March 2002 are allowed to choose either the old or the new policy. Shareable leave can be taken by either the fathers or the mothers, but not simultaneously, and in practice is mostly taken by the mothers. Full benefit compensation provides 90 percent of previous pay.

Appendix Table 2: Pre-birth characteristics before and after the 1984 extension

Characteristics in 1982	Before (se)	After (se)	Difference (p-value)
Mother's annual earnings (DKK)	197,110.2	195,644.6	1,465.7
	(951.2)	(893.4)	(0.261)
Father's annual earnings (DKK)	246,680.2	246,410.0	270.2
	(1590.1)	(1475.7)	(0.901)
Age of mother	26.24	26.08	0.166
	(0.05)	(0.0455)	(0.014)
Age of father	28.57	28.53	0.0385
	(0.069)	(0.063)	(0.680)
Mother with further education (%)	0.260	0.253	0.00768
	(0.005)	(0.0045)	(0.242)
Father with further education (%)	0.202	0.199	0.00356
	(0.004)	(0.004)	(0.554)
Mother's years of experience	5.41	5.379	0.0313
	(0.0385)	(0.0349)	(0.545)
Father's years of experience	6.901	6.929	-0.0279
	(0.043)	(0.040)	(0.635)
Number of observations	8352	9397	17749

Note: The table presents summary statistics of the characteristics in 1982 for two groups of mothers, those whose childbirths were between 1 January 1984 and 24 March 1984, and between 25 March 1984 and 18 June 1984. All characteristics are measured in 1982 for both the before and after groups. Annual earnings are measured in DKK in 2010. Further education is defined as having at least two years of schooling after high school/vocational level.
For each sample, the means are presented together with standard errors in parentheses. The last column presents differences between the before and after groups together with p-values in parentheses.

Appendix Table 3: Pre-birth characteristics before and after the 2002 extension

Characteristics in 2000	Before (se)	After (se)	Difference (p-value)
Mother's annual earnings (DKK)	225,309.2	222,537.8	2,771.4
	(1188.9)	(1137.2)	(0.090)
Father's annual earnings (DKK)	312,141.0	311,919.3	221.7
	(2075.8)	(1823.6)	(0.936)
Age of mother	29.13	29.04	0.0872
	(0.045)	(0.043)	(0.160)
Age of father	30.16	30.81	-0.650
	(0.082)	(0.062)	(0.000)
Mother with further education (%)	0.363	0.355	0.00725
	(0.0047)	(0.0045)	(0.262)
Father with further education (%)	0.284	0.288	-0.00427
	(0.0044)	(0.004)	(0.484)
Mother's years of experience	6.582	6.550	0.0319
	(0.0417)	(0.0396)	(0.579)
Father's years of experience	9.467	9.360	0.106
	(0.0537)	(0.0500)	(0.146)
Number of observations	10593	11442	22035

Note: The table presents summary statistics of the characteristics in 1982 for two groups of mothers, those whose childbirths were between 8 October 2001 and 31 December 2001, and between 1 January 2002 and 27 March 2002. All characteristics are measured in 2000 for both the before and after groups. Annual earnings are measured in DKK in 2010. Further education is defined as having at least two years of schooling after high school/vocational level. For each sample, the means are presented together with standard errors in parentheses. The last column presents differences between the before and after groups together with p-values in parentheses.