Parental Leave: Economic Incentives and Cultural Change*

James Albrecht
Georgetown Univ., IZA

Per-Anders Edin
Uppsala Univ.

Raquel Fernández
NYU, NBER, CEPR, IZA

Jiwon Lee
NYU

Peter Thoursie
Stockholm Univ.

Susan Vroman
Georgetown Univ., IZA

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Abstract

In 2002, Sweden reformed its parental leave system by adding a second "daddy month," i.e., a second month of pay-related parental leave reserved exclusively for each parent. In addition to giving fathers an economic incentive to take more leave, this change had an effect on cultural norms. We develop and estimate a model of the household in which preferences towards leave depend on the behavior of one's peers and use it to quantify the magnitudes of the economic-incentive effects as well as the evolving norms. We find that endogenously evolving cultural norms play a major role. We use our model to evaluate the effects of several potential policy changes including decreasing the cost of child care and giving each parent a non-transferable endowment of parental leave and conclude that only the latter would have a significant effect on the share of parental leave taken by men.

Keywords: Parental leave, gender equality, childcare, culture

JEL codes: D10, J16, Z10, Z18

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1 Introduction

Increasing gender equality is a central goal of many governments, multilateral institutions, NGOs, and activists. In this area, Sweden (along with Iceland, Norway, Finland, and New Zealand) excels, ranking in the top 5 countries in terms of its gender gap index in 2024.[1] The uptake of parental leave, however, as in all other countries in which universal parental leave is offered, remains very unequally distributed between women and men.[2] With the explicit goal of increasing gender equality, Sweden pioneered the introduction of earmarked, non-transferable, parental leave time (daddy months) in order to incentivize men to take a greater share of parental leave. This innovation – first introduced in 1995 by reserving one month of leave for each parent – is widely agreed to have increased men's share of parental leave.[3] In 1989, men took 6.9 percent of parental leave and by 1999 they were taking 10.6 percent. This reform was followed by the introduction of an additional daddy month in 2002 and a third one in 2016.[4]

The objective of our paper is to understand the role of the various potential determinants of parental leave uptake and how future policies may affect its gender distribution by using the second daddy month reform in Sweden. There are several factors that influence how parental leave is shared between parents. First, there are direct economic considerations as the loss of labor income is not fully compensated during parental leave. In addition, while childcare is very inexpensive in Sweden once a child turns one year old, that is not the case for a younger child for which parents must turn to the private market. This makes it very expensive for both parents to take short parental leaves that total less than a year. Second, there are penalties, both economic and social, associated with different leave times. In the economic sphere, individuals may face penalties in terms of their future wages and promotion prospects by taking "excessive" leave. From an employer's perspective,

[1] The Global Gender Gap Index benchmarks the current state of gender parity across four dimensions: economic participation and opportunity, educational attainment, health and survival, and political empowerment. See World Economic Forum (2024). Since the index's inception in 2006, Sweden has always ranked among the top 5 nations.

[2] According to Försäkringskassan (the Swedish Social Insurance Agency) at the end of 2022, men took 30% of parental allowance and women 70%. This is far from equal, of course, but a large increase over the 0.5% taken by men in 1974. See Försäkringskassann (2024).

[3] See, e.g., Ekberg, Eriksson and Friebel (2013) and Avdic and Karimi (2018).

[4] See Duvander and Cedstrand (2022) for an excellent review of the history of parental leave in Sweden.

an employee's length of leave time may signal something about their dedication to their firm or their degree of ambition. In the social sphere, people are concerned with how the length of their parental leave will be perceived and judged by others, independently of any material consequence. This includes the opinions of coworkers, friends and relatives, and even strangers in a playground, grocery store, or childcare center. Third, and related to the second, there is an idiosyncratic element governing how much a particular individual enjoys staying at home with a baby. This idiosyncratic factor, however, may well itself depend on social norms/culture as individual beliefs regarding what is "right" influence how content they are, and these beliefs are themselves dependent on social beliefs and the behavior of peers.

To understand and quantify how the forces above interact, we develop a simple choice model of a unitary household in which parents jointly decide how much parental leave each should take. Households, divided into four types by the education level of each parent (distinguishing between those with at least 3 years of university versus less), care about consumption (and hence income both during the parental leave period and after) and the amount of leave time each spends with the child. While all households are assumed to have the same utility function over consumption, an individual's preferences over parental leave time is not a pure primitive but rather depends also on the behavior of an individual's peers. In particular, we assume that an individual's enjoyment of parental leave time depends both on its length and on an idiosyncratic element that is drawn from a distribution whose mean is itself a function of peers' behavior. Peers consist of individuals of the same gender and, as will be made clear further on, also of the same household type.

We study the 2002 reform that introduced the second daddy month in Sweden and estimate our model using Swedish administrative data primarily from two time periods: the pre-reform period of 1998-2001 and the post two-daddy-month reform period of 2008-2011. We take (pre-birth) wages in each period as given and estimate, for men and women separately by education, the future wage penalty associated with different lengths of parental leave. While unobservable characteristics may influence this choice, these concerns are to some extent allayed by using individual fixed effects. We find significant changes in wage penalties between the two time periods which we then use as inputs to the model.

We ask our model to match key moments in the data, by gender and household type in *both* periods. This implies that the parameters governing the mean of the idiosyncratic draw of preferences influencing enjoyment of leave time, for example, are disciplined by the requirement of matching moments of the leave distribution both before and after the reform. Overall, given that we have four household types and two time periods, the model has 18 internally estimated parameters which are identified using the simulated method of moments by 56 moments in the data. We show that the model does a very good job in replicating the targeted moments. We also perform a regression discontinuity analysis to examine the immediate effects of the reform and find that the model does a reasonable job in generating quantitatively similar effects.

We use the estimated model to analyze how much of the changes in parental leave behavior over the two time periods is due to the changed opportunity set (i.e., the reform itself, independently of any other changes), income-related changes (wage parameters, wage penalties, and childcare costs), and the endogenous change in social norms. We find that the reform on its own is able to account for a small but not insignificant fraction of the changes in parental leave uptake. Income-related changes (in addition to the reform, but still keeping preferences invariant), also change behavior by a modest but significant amount. Keeping the economic environment at its pre-reform parameters, on the other hand, and allowing social norms to change in response to the reform has a significantly larger effect, substantially increasing men's parental leave and decreasing that of women. Lastly, when the income-related changes are introduced in addition to allowing preferences to change endogenously, the additional quantitative contribution tends to be even larger. This reflects the extra kick that comes from preferences further changing in response to the income-related changes in parental leave behavior.

We then use the model to investigate the role of the gender wage gap and conduct three counterfactual policies, all aimed at increasing gender equality in parental leave. The first policy subsidizes the cost of childcare during the first year, equating it to the low cost of publicly provided childcare that is only available once the child turns one. The second policy (the "endowment policy") is similar to current policy in Iceland and Finland. It endows each parent with six months of non-transferable leave (and one month that can be

transferred freely). The last policy (the "equal share policy") is significantly stricter in its approach to gender equality: with the exception of one month that can be freely allocated, a parent's length of (paid) leave is not allowed to exceed that of the other parent. We find that decreasing the cost of childcare has very little effect on the share of parental leave taken by men. The last two policies, on the other hand, significantly increase men's share of parental leave.

Our paper is related to a few areas of research. First, there is a small literature that studies the determinants of the within-couple allocation of parental leave. Jørgensen and Søgaard (2024) estimate parents' willingness to pay for parental leave by exploiting the fact that the Danish parental leave system creates important kinks in the household budget set. This is a consequence of the combination of low public benefits with private provision by employers which vary in length, leading to significant bunching at various kink points. The authors combine a semi-structural estimation approach with features of the nonlinear budget set literature (Saez (2010)) to estimate the joint preferences of the household and to study the effect of various counterfactual reforms. Our paper also focuses on household preferences but does not build on the public economics literature. Instead it complements that literature by explicitly modeling costs such as future income penalties associated with leaves of different lengths, heterogeneity in preferences by education and, most importantly, modeling culture and allowing it to endogenously change in response to reforms. A recent working paper by Linderoth (2024) also uses bunching at kink points, but assumes that Swedish households' preferences over the division of parental leave are reference dependent, with the daddy-month quotas as reference points. The paper concludes that reference dependence is more important than financial incentives in determining fathers' leave-taking behavior. Our findings echo this conclusion in the sense that purely economic considerations have only small effects at the margin, but in our model culture is determined by what other people do rather than by government policy directly and hence evolves endogenously.[5]

There is also a large and growing literature on the effects of parental leave. These

[5] A different possibility – that leave shares are determined by physiological demands – is ruled out by the findings of Moberg and van der Vleuten (2022). They use administrative data to compare the pattern of parental leave uptake among biological parents to that among adoptive parents and conclude that the patterns observed are not due to the biological demands of motherhood, but rather are more likely explained by financial incentives and/or norms.

papers have focused on a diverse set of issues ranging from the consequences of parental leave for maternal welfare and children's health and human capital (e.g., Cools, Fiva and Kirkebøen (2015), Carneiro, Løken and Salvanes (2015), Dustmann and Schönberg (2012), and Rossin-Slater (2018)), fertility (e.g., Lalive and Zweimüller (2009) and Farré and González (2019)), and labor supply at home and in the labor market (e.g., Dahl et al. (2016), Duvander and Johansson (2019), and Schönberg and Ludsteck (2014)), among others. Excellent surveys of this literature can be found in Olivetti and Petrongolo (2017) and Canaan et al. (2022).[6] On the structural side, Erosa, Fuster and Restuccia (2010) use a general equilibrium search model to study how parental leave affects fertility and labor market decisions and Yamaguchi (2019) estimates a dynamic discrete choice model of female employment and fertility to evaluate how different lengths of job protection and paid family leave might affect women's career and fertility choices.

There is a small number of papers showing the importance of peer effects for parental leave decisions. Dahl, Løken and Mogstad (2014) study a 1993 Norwegian reform that earmarked an extra month of leave for fathers exclusively. Using a regression discontinuity design, they show that this reform significantly increased the proportion of fathers taking paternity leave immediately after the reform. More importantly from our perspective, the authors show that a man who had a child in a window right after the reform increased the subsequent take-up of paternity leave by his coworkers and brothers relative to their counterparts who had a peer who became a father in the window prior to the reform. Welteke and Wrohlich (2019) examine peer effects stemming from a German reform in 2007 that changed parental leave payments from means tested to universal and also changed the length of time these were paid. Exploiting whether a woman had peers who were affected by the reform, they found that a woman's choice of length of parental leave was affected by her coworkers' decisions. Carlsson and Reshid (2022) use a "peers of peers" strategy by instrumenting co-workers' leave uptake in Sweden with the leave of their siblings and cousins. They too find substantial peer effects for both women and men. Lastly, a recent paper by Dottori, Modena and Tanzi (2024) exploits an Italian reform in 2015 that increased the flexibility of paid parental leave time from requiring its use by age four of the child to

[6] See Ruhm (1998) for an early survey.

age six. The authors find that the reform increased the takeup of parental leave for mothers especially if the share of post-reform peers taking leave was larger.

Our paper contributes to the literature on cultural change, especially regarding gender roles.[7] Kotsadam and Finseraas (2011) show that the Norwegian daddy quota led to greater sharing of household labor (laundry) and fewer conflicts over these chores. Using the introduction in 2007 of paternity leave that increased by 13 days the number of fully compensated leave days in Spain, Farré and González (2019) show that fathers become more involved in childcare. Intriguingly, there is also evidence that the effects of paternity leaves spill over to the children's generation. Using a 2010 Norwegian survey of high school students, Kotsadam and Finseraas (2013) show that girls (but not boys) born right after the 1993 Norwegian paternity leave reform were less likely to report doing housework compared to those born right before. Farré et al. (2023) show that the 2007 Spanish paternity leave reform affected the gender norms of children born right after the reform. These were more likely to support more gender-equal attitudes and both genders were more likely to engage in "counter-stereotypical" household tasks.

Our paper's main novel contribution is to model cultural change and to quantify how it matters for outcomes. It complements the causal empirical literature by allowing one to disentangle the relative importance of different mechanisms and to study the potential effects of alternative reforms, both of which would be difficult with a purely reduced-form approach. In this sense, it is closest to Fernández (2013) which estimates a quantitative model of married women's labor force participation over 120 years in the US in which both economic factors – men's and women's wages – and cultural beliefs play a role in determining women's work decisions. That paper, however, models cultural change as the (Bayesian) evolution of beliefs generated by women learning about the true cost of working through observing the decisions of prior cohorts. In our model, on the other hand, social norms are captured by the mean of the distribution governing idiosyncratic preferences towards spending time on parental leave. These evolve endogenously as a result of prior decisions made by individuals who are similar in various dimensions (gender and household type).

Our paper is organized as follows. The next section briefly discusses the Swedish parental

[7] See Fernández (2025) for a review of the literature on cultural change.

leave system and the time periods of the analysis. Section 3 presents the datasets and the sample construction. Section 4 examines the distribution of parental leave takeup and estimates the future wage penalties. Section 5 presents the model and section 6 its estimation. Section 7 disentangles the contributions of the reform, of the changed wage parameters, and of changes in culture or social norms to the takeup of parental leave by gender and household type. Section 8 re-examines the effect of the reform alone by contrasting its results with a regression discontinuity analysis. Section 9 studies several potential alternative policies and the role of the gender wage gap. Section 10 concludes.

2 Some Preliminaries

2.1 The Parental Leave (PL) System

Sweden has been a leader in parental leave. In 1974 it became the first country to extend paid parental leave equally to fathers.[8] This reform granted 6 months of earnings-related benefits for parental leave at a 90% replacement rate up to a fairly generous ceiling. Mothers and fathers were free to split these 6 months as they pleased. The system was extended several times and, by 1994, 12 months of earnings-related benefits were available plus an additional 3 months at a very low flat rate. This was complemented by a public daycare system that became available once a child turned one year old, and most parents avail themselves of this option.[9] In addition, parental leave could be taken on a part-time basis and could also be taken later (up to when the child was 8 years old).[10] Furthermore, a parent's job was protected for up to 18 months.[11]

An important change in Sweden's parental leave system was the introduction of months reserved for each parent – an innovation that Sweden spearheaded and that later spread to other countries, particularly in the European Union. The first "daddy month" was introduced in 1995. In principle, the reform allocated one of the earnings-related months

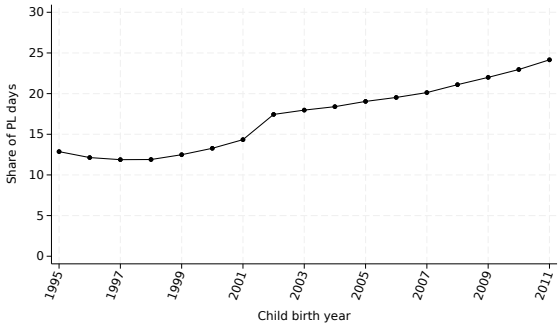
[8] Before 1974, there was a less generous leave policy that was reserved solely for mothers.

[9] For example, in 2004, 74 percent of all children aged 1-3 and 96 percent of children 4-5 were registered in childcare (Skolverket (2005)).

[10] More than 90% of all leave days are taken in the first two years, however (Ekberg, Eriksson and Friebel (2013)).

[11] See Duvander and Cedstrand (2022) for more detail.

Figure 1: Evolution of Fathers' Share of PL Days



Note: Fathers' share is calculated using gross PL days (defined below) taken within the first two years after childbirth. These shares are derived from data provided by the Social Insurance Agency and include all births.

of benefits to each of the parents. The remaining 10 months could be used as the parents wanted. In practice, since most parental leave was taken by mothers, this meant that the father needed to take at least one month of parental leave or it would be lost; hence the name "daddy month." In 2002, a second month of benefits earmarked to each parent was introduced (the second daddy month). The maximum duration of earnings-related parental leave was simultaneously increased by one month from 12 to 13 months meaning that, in practice, women could still take 11 months of earnings-related leave as before. We study this reform.[12] Figure 1 shows the evolution of father's share of parental leave days over time. Note the increased slope between 2001 and 2002 when the second daddy month was introduced.

Since 1998, the earnings-related replacement rate has been constant at 80%. Earnings related payments are subject to a ceiling, and this ceiling eroded to some extent over time until it was raised again in 2006. On the other hand, many unions have collective agreements that top up the benefits that their workers receive. On net, our judgment is that the effect of the top-ups exceeds that of the erosion of the ceiling. In our analysis, we therefore assume an effective earnings-related replacement rate of 85%. The flat rate that is available once earnings-related payments are exhausted was sufficiently low that we simply set it to zero in our model.[13]

[12] In 2016, a third month of earmarked benefits was added to the system.

[13] The flat rate was 60 SEK per day (approximately 9 USD per day) from 1998 to 2001, the earlier period that we analyze, and was 180 SEK per day (approximately 27 USD per day) from 2008 to 2011, the later period that we analyze. This nominal increase corresponds to more than a doubling of the real rate.

2.2 Time Periods of Analysis

To understand the role of economic and societal forces in changing household behavior, we analyze the response to the 2002 reform. This reform introduced a second daddy month and also lengthened the earnings-related parental leave period to 13 months. We examine this reform rather than the first reform of 1995 for several reasons: i. an increase in the intensive margin of reserved months is interesting from a cultural perspective; ii. several papers have focused on the 1995 reform already and, most importantly, iii. the data provided by the Social Insurance Agency is incomplete for 1994, the first year in which PL data from the agency is available. While PL days taken in 1993 and extending into 1994 are recorded, those that ended before 1994 are missing. Since parents can take parental leave up to 60 days before childbirth, a significant number of days may be missing for children born in early 1994.[14]

Our estimation is disciplined by focusing on two four-year time periods: one before the reform and one after. The "pre-reform period" is from 1998 to 2001. During this period there was only one daddy month and 12 total months of earnings-related parental leave. We chose the years to allow for the maximum adjustment to the one daddy month reform without overlapping with the 2002 reform. The "post-reform period" is from 2008 to 2011. We chose this period to be six years after the introduction of the second daddy month to allow time for changes in behavior, especially those driven by changes in cultural attitudes and wage penalties.[15]

During our period of analysis, there were of course other policy changes in Sweden related to parental leave. These include a child care reform in 2002 (Maxtaxan), a gender equality bonus implemented in 2008, and the introduction of "double-days," allowing both parents to take parental leave simultaneously, in 2011. Lundin et al. (2008) found that the 2002 child care reform, which reduced child care fees for some parents and extended entitlement to child care for parents who are unemployed or on parental leave, did not affect labor supply. The "gender equality bonus" was a tax incentive for parents who split

[14] This issue is less problematic when applying a regression discontinuity approach to analyze the 1995 reform, as in Avdic & Karimi (2018), since their analysis focuses on PL days around the January 1, 1995 cutoff.

[15] As noted in the data section, missing parental leave data in 2013 means that starting the second period later would be problematic.

their parental leave equally. Duvander and Johansson (2012) concluded that it did not change parental leave decisions perhaps because it was complicated and received less public attention. The double-days reform only affected the last cohort of children in our sample and had a low take up. Given these findings, we think that these reforms had negligible effects on parental leave take up.

Between 1995 and 2001, men's share of parental leave evolved slowly from 12% to 14%. By 2009, men were on average taking 20% of parental leave time. It is important to take into account that the parental leave behavior of both men and women differs by education. The proportion of university-educated men (those with at least 3 years of university) taking more than 2 months of parental leave was 32% in 1995 while that of non-university men was 21%. By 2009, these figures were 67% for the university men and 44% for the non-university men. Similarly, the proportion of women taking strictly less than 11 months of parental leave is significantly higher for university-educated women. In 1995, this figure was 30% for the university-educated women and 16% for the non-university women. By 2009, these figures were 40% and 26%.[16] Attitudes also differ significantly with women being significantly more in favor of women working full time when they are parents and more educated individuals being more favorable than those with less education. These observations suggest that education is an important factor to take into account in our analysis.

3 Data and Sample Construction

This section describes the main Swedish administrative registers from which we derive our data and discusses the construction of the samples that are used in our analysis.

3.1 Datasets

The Parental Leave Registers constructed by the Social Insurance Agency contain information on the number of paid parental leave days taken by each person.[17] We use the start and end dates for parental leave spells to construct a measure of the months an individual spent

[16] The figures in this paragraph are based on the authors' calculations using the couples dataset defined in the next section.

[17] The parental leave data cover the period from 1994 to the first quarter of 2022. There is no data, however, from October to December 2013.

on parental leave in the first 16 months after the birth of their child. A spell represents a period of time that a parent informs the Social Insurance Agency that they will be away from work on parental leave. We use *gross* months as our measure, meaning that the person may be receiving only partial parental leave payments and may either be working part time or simply choosing to smooth payments and not working. The data does not allow us to differentiate between these possibilities.[18] For each parent, our parental leave variable is the sum of the length of all spells taken in the first 16 months. Note that mothers-to-be have the option to start taking paid leave 2 months prior to birth. This means that some women are recorded as taking more than 16 months of leave. In our analysis, we group individuals who are recorded as taking 17 months or more and code them as 17 months.[19] Restricting our window to 16 months avoids confounding parental leave associated with the first birth with parental leave associated with subsequent births.[20] Furthermore, parental leave data are unavailable in the October-December 2013 window. Allowing for more than 16 months of parental leave after childbirth would limit the length of the post-period.

The Multi-Generation Register records all births by month and year. We use this dataset to link children to their biological parents. To obtain other characteristics of parents, we use LISA.[21]. This administrative database includes information on parents' education levels and also records individuals' annual earnings as well as standard background characteristics such as a parent's age, country of origin, and region of residence. We use these data to classify individuals by education (non-university versus university) and, as described below, to limit our sample to Swedish-born parents.

Finally, we link the data described above to the Wage Structure Statistics database to obtain full-time-equivalent monthly earnings. These are based on employer reports of individual earnings and contracted hours during a survey month (typically September). We use full-time-equivalent monthly earnings as an individual's monthly wage, even if the individual did not work full time. Of course, these data are only available for individuals who

[18]The data only includes annual earnings and does not specify when income was earned. Additionally, there are no labor supply measures that indicate how many hours the individual worked over the year.

[19]The percentage of women who take more than 17 months is always significantly under 1%.

[20]Only 4 percent of couples have an additional child within 16 months.

[21]The Longitudinell Integrationsdatabas för Sjukförsäkrings- och Arbetsmarknadsstudier or, in English, the Longitudinal Integration Database for Sickness Insurance and Labor Market Studies

had strictly positive working hours during the survey month. It is important to note that although these monthly wage data are available for all public-sector employees with positive hours in the survey month, this is not the case for the private sector. For the latter, wages are available for workers in firms with 500 or more employees and for a stratified sample (based on industry and firm size) of workers in smaller firms. As a result, wage data are available for roughly 50 percent of private-sector workers.

3.2 Sample Construction

We restrict the sample in several ways. First, we include only first births for Swedish-born women and men where we observe both parents in our data. Focusing on first births allows us to avoid issues related to the relative timing of prior births which may affect the uptake of parental leave. Restricting the sample to Swedish-born parents eliminates issues concerning when immigrant parents came to Sweden and whether their use of the parental leave program is affected by their immigrant status. Moreover, the cultural attitudes of immigrants is very likely to differ from those of native Swedes.[22] Second, we limit our sample to heterosexual couples between the ages of 20 and 60 at the time of childbirth as same-sex couples are likely to have different social norms regarding parental leave.

As noted previously, the second daddy month was introduced in January 2002 and we therefore restrict our attention to births in the years 1998 to 2001 – the pre-reform period – and in the post-reform period of 2008 to 2011. Starting the post-reform period with 2008 allows six years for the adjustment of wage penalties and cultural change in reaction to this policy change. In total, this leaves us with 226,369 childbirths (102,547 in the pre-period and 123,822 in the post-period).

The Wage Sample

The wage penalties associated with different lengths of parental leave are an important part of our analysis. As described in the next section, to estimate these wage penalties, we use data on individual monthly wages, i.e., full-time monthly equivalent earnings, three years after childbirth as well as one year prior to the birth.

[22] See the discussion in Tervola, Duvander and Mussino (2017).

Eliminating parents with missing wage observations in those years substantially reduces our sample, leaving us with 68,354 mothers (29,178 in the pre-period and 39,176 in the post-period) and 64,714 fathers (28,285 in the pre-period and 36,429 in the post-period).[23] Note that we include individuals even if we do not have wage data for their partners. We refer to this as the "wage sample."

The Couples Sample

We take the couple as our unit of observation for parental leave, as a leave decision made by one parent necessarily affects the other. This requires us to further restrict our sample to observations for which we have wages for *both* parents, reducing the sample to 9,926 couples in the pre-period and 13,796 couples in the post-period.

As discussed previously, education appears to be an important determinant of the length of parental leave. We classify individuals into university (u) versus non-university (n) categories according to whether they have at least three years of university studies versus less. This partitions our couples into four household types, denoted by nn, un, nu, and uu, where the first entry indicates the father's education and the second the mother's.[24] Table A2 presents summary statistics for the wage sample and the couples sample discussed in this section.

4 Parental Leave and Future Wage Penalties

In this section, we turn to a description of some of the key features of our data. We provide statistics on parental leave length by gender and household type and we estimate, by gender and education, wage penalties associated with different lengths of parental leave.

[23] This is not a result of Swedish women not working after having children. Eliminating those who are on parental leave with another child three years later, over 90% of mothers were back in the workforce.

[24] Table A1 in the Appendix gives the distribution of household types in the couples samples by period. It is worth emphasizing that changes in the distribution of household types does not affect our analysis as all the moments used for estimation are *by household type*.

4.1 Parental Leave

In this subsection, we show the distributions of parental leave across mothers and fathers in the pre- and post- periods. As we noted in the previous section, we restrict our attention to parental leave spells that occur in the first 16 months after birth.

To receive parental leave benefits, a parent needs to inform the Social Insurance Board. We measure parental leave by the length of the periods that a parent informs the Social Insurance Board that they will be on parental leave. Although parents may take less than full benefits in these periods, they are not allowed to both working full time during these periods. We measure parental leave in months so that a parent who does not take any parental leave has leave equal to 0, while leave equal to t months indicates that the parent took more than t-1 months, but no more than t months, i.e., t months indicates PL $\in (t-1, t]$.

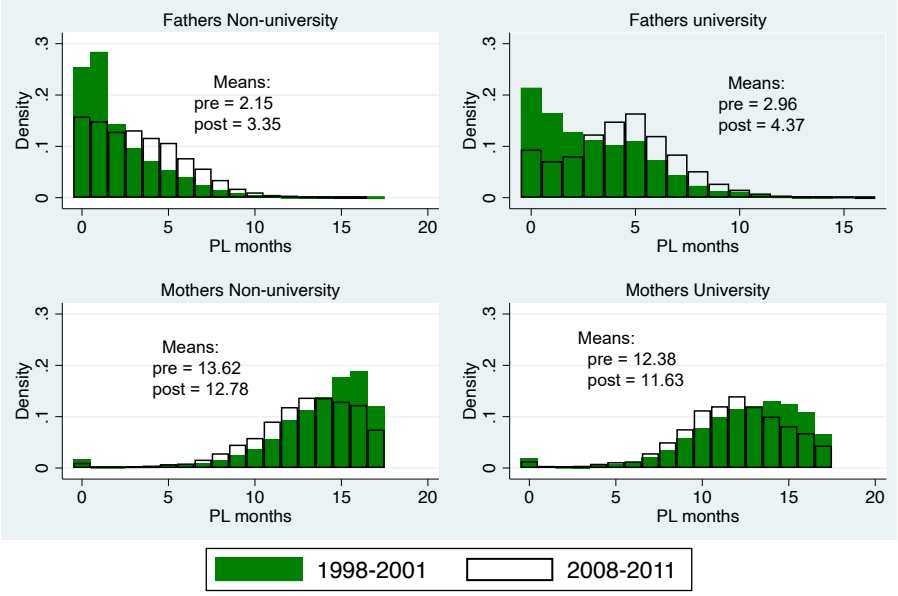
Figure 2 displays the distribution over months of parental leave by gender and education in the pre-reform (green) and post-reform (white/transparent) periods. Note that the change in men's parental leave distribution occurred over the entire range of the distribution, as did women's. It is also worth noting that a large percentage of men, over 20% regardless of education, were taking zero months of parental leave prior to the 2002 reform. Women's parental leave distributions show a significant reduction in long duration parental leave, i.e., leaves of 15 or more months.

Lastly, we can take the household as our unit of observation and, by averaging over the share of leave taken by the father in each couple, show how the distribution of the father's share of parental leave changed between the two time periods. This is portrayed in Figure 3, distinguishing here only by the man's education category. Once again, we see that men's share of parental leave shifted everywhere to the right in the post-reform period.

4.2 Future Wage Penalties

In this section we estimate the wage penalties associated with different durations of parental leave. These wage penalties are potentially an important input into the couple's decision about how much leave to take. While there is now a large literature that examines the effect of parenthood on gender wage and earnings gaps (see Albrecht et al. (1999), Angelov,

Figure 2: Parental Leave Distributions by Education and Sex



Note: Based on authors' calculations using the couples sample.

Johansson and Lindahl (2016), and Kleven, Landais and Søgaard (2019)), here our focus is on the effect of parental leave duration on the within-gender and education distribution of wage growth.

To estimate wage penalties, we use the full wage sample described earlier, i.e., we do not restrict the analysis to the couples subsample. All wages are expressed in 2014 SEK. The underlying assumption is that an individual's wage penalty is driven by their gender and education rather than household composition.

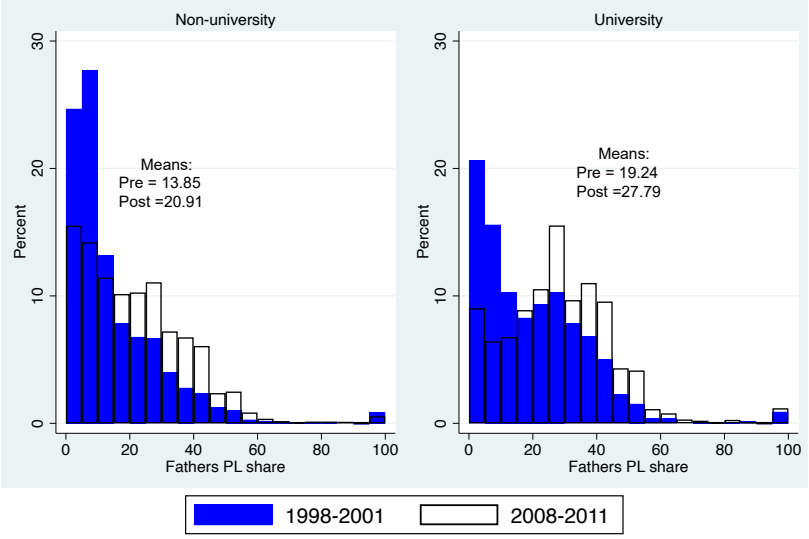
We examine wages both three years and one year prior to the birth of the first child. Accordingly, we specify the following wage equations for these two time points:

$$\log(w_{i,t+3}) = \alpha + \beta_1 plcat_i + \beta_{2,t+3} agecat_i + \beta_3 \text{YoB}_i + \gamma_i + \lambda_{t+3} + \epsilon_{i,t+3} \quad (1)$$

$$\log(w_{i,t-1}) = \alpha + \beta_{2,t-1} agecat_i + \beta_3 \text{YoB}_i + \gamma_i + \lambda_{t-1} + \epsilon_{i,t-1} \quad (2)$$

where $\log(w_{i,t})$ denotes the logarithm of the wage of individual i in year t. The variable

Figure 3: Fathers' share of parental leave



Note: Father's education is used to divide sample here. Calculations use the couples sample.

$plcat_i$ is a categorical variable (defined below) indicating the number of months of parental leave taken by individual i within 16 months following childbirth. This variable is excluded from Equation (2), as parental leave cannot be taken in years prior to childbirth. The term $agecat_i$ is a vector of age-category dummies based on the individual's age at the time of childbirth, with the associated coefficient β_2 allowed to vary over time. YoB_i denotes a vector of birth-year fixed effects, λ_t is a year fixed effect, and $\epsilon_{i,t}$ is the error term. Wage equations are estimated separately for each gender-education group and for each period (pre- and post-reform).

A potential concern is selection into different durations of parental leave. For instance, individuals performing poorly at work before the birth of their first child may be more inclined to take longer leave, and subsequently continue to experience low wages—potentially due to the same underlying factors that influenced their pre-birth labor market position. Estimating the model in first differences addresses this concern by eliminating time-invariant individual characteristics, both observed and unobserved. However, we do not interpret the resulting estimates causally, as selection on time-varying unobservables remains unaddressed.

We therefore estimate the following differenced equation:

$$\log(w_{i,t+3}) - \log(w_{i,t-1}) = \beta_1 plcat_i + \beta_2^* agecat_i + \lambda_t^* + \epsilon_i^* \tag{3}$$

which yields an estimate of the wage penalty β_1 three years after childbirth. Here, $\beta_2^* = \beta_{2,t+3} - \beta_{2,t-1}$ and λ_t^* captures the difference in year fixed effects between $t+3$ and $t-1$. In practice, these effects are absorbed by inclusion of cohort fixed effects, as YoB_i and year fixed effects are not separately identified in the differenced model. Finally, ϵ_i^* denotes the difference in the individual error terms across the two time points.

Our estimate of equation 3 groups men with 9 months or more of parental leave into one category of 9 months and women who took 10 months or less into a category of 10 months as there are too few individuals in the outlier months. Plots of the estimated coefficients are given in Appendix 5. Table 1 reports these wage penalties averaged within intervals of months over which the coefficients are similar (but not all individually significant). The omitted category for men (in both periods and for both levels of education) is zero months of parental leave, and the omitted category for women (again, in both periods and for both levels of education) is 0 to 10 months of parental leave. The penalties reported here are used as an input in the estimation of the parental leave choice model.

Table 1 shows several interesting patterns. For non-university men, there are penalties associated with taking any amount of parental leave. In the pre-period, these men suffer a wage penalty ranging from about 2% for 1 to 4 months of parental leave to 3.8% for 9 months and above. In the post-period, the penalty is smaller than in the pre-period and also rises with parental leave uptake, going from about 1.4% for 1 to 7 months of parental leave to 2.5% for more than 8 months. In both the pre- and post-periods, university men can take several months of leave without incurring any penalty (3 months in the pre-period and 4 months in the post-period) but there is a substantial penalty of 2.9 % associated with taking 4 to 8 months and 5.9% with taking 9 months or more in the pre-period. In the post-period, the penalties are smaller ranging from 1.7% for 5 to 8 months and 3.6% for 9 months and more.

Non-university women in the pre-period suffer no wage penalties for taking up to 11

months of parental leave. Those who take more than 11 months suffer relatively small wage penalties ranging from 1.3% for 12 to 14 months to 2.4% for 17 months. In the post-period, these penalties are even smaller. There is no penalty associated with 12 months or less of leave and the penalty for taking 13 months or more is about 1.2%. Finally, among university women, in the pre-period, there are penalties associated with taking 11 months of leave or more. These range from 1.6% for taking 11 to 12 months of leave to 4.6% for taking 17 months of leave. In the post-period, there are also penalties for taking 11 months or more, but these are uniformly lower than those of the pre-period.

Overall, on average men's parental leave penalties decreased between the pre and post-periods as did those of women. While we feel fairly comfortable interpreting the findings for men as reflecting a reduced penalty given the change in norms and hence how employers view a signal of a lengthier parental leave, it is trickier to interpret the results for women. The evidence presented in Kleven, Landais and Søgaard (2019) suggests that Danish women who work after the birth of their first child suffer wage penalties because they are more likely to switch to working in the public sector or in a firm with a female manager (both more likely to be family friendly) as well as to reduce their hours and their occupational rank. In Sweden, few women switched to working in the public sector in either period and a greater percentage of them chose to work full time or close to it in the post-reform period. Furthermore, a smaller percentage switched employers in the post period.[25] Thus, the reduced parental leave duration for women may reflect a greater attachment to their careers – also a change in norms – rather than employers caring less about the duration of women's leave.[26] We have also re-estimated the model attaching zero wage penalty to any amount of leave chosen by women both in the pre and post periods, and our main results

[25] In the pre-reform period, 4% switched from private to public-sector jobs whereas 7% switched in the opposite direction. The corresponding figures for the post-reform period are 3% and 5%. Of the women working three years after the birth of their child, 67% were working full time in the pre-reform period whereas 78% were in the post period, where here we denote by "full time" as working more than 75% of standard "full time." Fewer women switched employers at t+3 in the post period as well: 36% in the pre-reform period versus 26% in the post-reform period. These last figures are derived from Statistics Sweden's Business Register, which includes information on all companies, government agencies, and organizations in Sweden. We match this data with the individual-level data used in our analyses.

[26] Tô (2018) constructs a model of the effects of parental leave on mothers' subsequent wages in which the decision to take fewer months than the maximum signals career commitment. Using Danish data, she finds that when the allowed maximum increases, mothers who forego some of the additional paid parental leave are rewarded with higher subsequent wages relative to those who take the full allowed amount.

(e.g. those of Section 7) are very similar.[27]

Table 1: Estimated Wage Penalties

	Men				Women			
	Non-Univ		Univ		Non-Univ		Univ	
Period	Intervals	Wage Penalty	Intervals	Wage Penalty	Intervals	Wage Penalty	Intervals	Wage Penalty
Pre	1 - 4	-0.0199	1 - 3	0	11	0	11 - 12	-0.0163
	5 - 8	-0.0308	4 - 8	-0.0294	12 - 14	-0.0128	13 - 15	-0.0285
	9 - 17	-0.0380	9 - 17	-0.0588	15 - 16	-0.0188	16	-0.0378
					17	-0.0241	17	-0.0458
Post	1 - 7	-0.0137	1 - 4	0	11 - 12	0	11	-0.0113
	8 - 17	-0.0249	5 - 8	-0.0173	13 - 17	-0.0118	12	-0.0159
			9 - 17	-0.0355			13	-0.0224
							14 - 15	-0.0271
							16 - 17	-0.0314

Note: The calculated wage penalties use the coefficients from the wage penalty regressions, as shown in equation 3. The wage penalties presented in this table represent averages of the individual monthly wage penalties ($\beta_j^{g,e}$) over intervals of months over which the coefficient values are similar. The omitted category for men is zero months of parental leave; for women it is 10 months or fewer. See Appendix Figure 5 for estimated coefficient plots.

5 A Simple Model of Parental Leave

As noted previously, the objective of the daddy month reform was to increase the amount of parental leave taken by fathers. There are several reasons this could happen. First, there is a direct incentive effect. The daddy month is an extra month which can be taken for close to full pay. If a father fails to take his daddy month(s) and that time is instead covered by the mother at the low flat rate, the couple is "leaving money on the table" in the sense that the father could have been home with the baby at a lower cost to current household income. Second, as shown in Section 4.2, men's wage penalty decreased between the pre- and post-reform periods, increasing men's incentive to take more parental leave. A third factor operates through changing social norms or culture. Men may be unwilling to take longer parental leave if other men are not doing so. As more men respond to the direct financial incentives associated with daddy months, other fathers may be more comfortable taking additional parental leave themselves.[28] This would also affect mothers' behavior as their desire to take more months of low-paid parental leave might change if their partners

[27] Results available from the authors.

[28] Dahl, Løken and Mogstad (2014) cite a Norwegian survey taken prior to the parental reform of 1995 as saying "Fathers are concerned that both employers and coworkers will perceive them as less invested in their careers if they exhibit a large commitment to family."

are now taking more leave. This would in turn also decrease the social pressure facing women to take lengthy spells of parental leave.

5.1 The Model

We next present a model that captures these incentives and concerns in a simple and transparent fashion. To this end, consider a unitary family of a woman and man who have just had their first child and need to decide how many months of parental leave they each wish to take over a period of length τ of which there is a maximum of T months of highly paid parental leave. These would be 12 months in the pre-reform period and 13 in the post-reform period. While technically there are also an additional 3 months available at a low flat pay, we will treat these months as paying zero as noted previously but include them of course as PL months in our calculations. Note that since we do not distinguish between partial and full benefits, we effectively allow 17 months as our window in which to observe parental leave, as described in Section 3.

Let t_m and t_f be the parental leave taken by fathers and mothers (measured in months) where m and f denote male and female, respectively. We assume that the household preferences can be modeled as:

$$V_{ij} = U(\tilde{y}_{ij}) + B \cdot U(y_{ij}) + \gamma_m^{ij} \cdot h_m^e(t_m) + \gamma_f^{ij} \cdot h_f^e(t_f) \tag{4}$$

where $B > 0$ and i, j denotes the household type (i.e., the education i of the man and j of the woman, $i, j \in \{n, u\}$).[29] The first term is the utility over the average monthly family income $\tilde{y}$ received during the period of potential parental leave τ, where

$$\tilde{y} = \tilde{y}_m(t_m, t_f) + \tilde{y}_f(t_m, t_f) - C(t_m + t_f)$$

is the average net household income over the 17 month period, i.e., it is the sum of both parents' labor and parental leave incomes net of the total cost of child care C during that

[29] In order to simplify notation, equation 4 has suppressed individual-specific subscripts. In particular, as will be made clear later, both income and γ are individual-specific and do not only depend on gender and household type.

period. The second term is the utility over average monthly *future* family income

$$y = y_m(t_m) + y_f(t_f)$$

which depends on each parent's parental leave time as these determine the future wage penalties associated with these decisions. As discussed previously, the length of parental leave each individual takes potentially affects future income as firms may interpret longer leave times as signaling something about the worker's type such as a lower commitment to work, to the firm, or to their career.

The third and fourth terms represent the utility that each parent obtains from the time spent with their child during parental leave. We model this as an idiosyncratic enjoyment of spending time with one's child, γ, and a function of the time spent $h(t)$.[30] These preferences are not pure primitives, however, as γ is also assumed to depend on the behavior of others.

In particular, we assume that an individual's γ is drawn from a normal distribution $\mathcal{N}(\mu_\gamma^{g,ij}, \sigma_\gamma^{g,e})$ whose mean evolves according to the following law of motion:

$$\mu_s^{g,ij} = a^{g,e} + b^{g,e} \cdot \bar{t}_{s-1}^{g,ij} \tag{5}$$

where $\bar{t}_{s-1}^{g,ij}$ denotes the mean parental leave taken in the previous period, $s-1$, by individuals of the same gender g and household type ij. Note that the parameters in equation 5, $a^{g,e}, b^{g,e}$ are allowed to depend on gender and education as suggested by the poll evidence; the standard deviation, $\sigma_\gamma^{g,e}$, is assumed to stay constant. This formulation captures the idea that how an individual feels about taking parental leave depends on the choices made by others. The restriction to caring about the choices of same-gender individuals is natural. Caring more about the choices made by those who are similar to one is also intuitive. One possibility would be to care only about the choices made by those with the same education. Here we go a step further and assume that the education of one's partner also matters. That is, women care about the choices made by women with their same education and whose

[30] We chose to model the utility from parental leave as separable in men's and women's time. An alternative would be to have the weighted sum of parental time matter, although it would require us to also estimate the weights attached to each time input. Furthermore, we would need to introduce two additional terms to capture social norms for each parent. We think that the present specification is more transparent and does not require us to model social norms separately from preferences.

partner has the same education as their partner, and similarly for men. This can reflect that individuals are influenced not only by their coworkers and friends, who are likely to be of a similar education as themselves, but also by those with whom they socialize, which is likely to depend as well on the education of their partner.

For simplicity, we assume that the mean of the γ distribution is a linear function of the mean time spent on parental leave in the previous period by those similar to oneself (as defined above), $\bar{t}_{s-1}^{g,ij}$. In equation 5, a fixes the intercept whereas the marginal response to changes in others' parental leave choices is given by b. Lastly, we assume last period's, rather than the current period's, mean parental leave influences μ. As shown by Dahl, Løken and Mogstad (2014) in the context of a parental leave reform in Norway, there is a "snowball" effect of past parental leave choices made by coworkers whose effects accumulate over time. While this does not matter in a steady state, it implies that studying the effect of a reform may require time, as social interactions gradually lead to a higher long-run equilibrium.[31] Furthermore, this formulation allows the model to avoid the usual multiplicity of equilibria that could arise if μ depended instead on the current mean level of parental leave.

5.2 The Maximization Problem

The couple chooses their parental leave durations, t_m and t_f, to maximize their welfare in equation 4 subject to the following constraints: (i) τ is the maximum amount of observed leave time and therefore, effectively, the amount of leave that an individual can take independently of whether it is paid or not. For the estimation, we take $\tau = 17$ months. As discussed previously in Section 3, we treat months as a discrete variable, so $\tau = 17$ implies $t_g \in \{0, 1, ..., 17\}$ for $g \in \{m, f\}$. (ii) Let T be the maximum possible number of highly paid months. If the couple jointly takes more than T months of leave and both partners take at least D, where the latter is the number of daddy months, parental leave is paid for T months and the remaining months of parental leave pay zero.[32] (iii) If a parent j takes fewer than D months, then the aggregate parental leave constraint at which pay becomes

[31] We will not estimate the transition path and therefore do not define what time length to assign to a period.

[32] As noted above, there are three months that are paid at a sufficiently low flat rate that we set to zero and we do not distinguish between full and partial benefits.

zero becomes $T - (D - t_j)$.

We assume a uniform monthly wage replacement rate of $\kappa = 0.85$ when a worker takes highly paid parental leave.[33] Furthermore, we assume that childcare costs are incurred whenever both parents are working. The monthly cost is C_r when the child is less than a year old and hence the parents are assumed to use private child care, and it is C_u when the child is a year or older and can use public child care, with $C_r > C_u$. We set C_r and C_u equal to 50% and 9.1%, respectively, of the average monthly wage of a non-college woman in the pre-period.The corresponding numbers for the post-period are 45% and 3.6%, respectively.[34] Appendix Section 10.3 derives how a household will divide the allocation of a given number of months between the two parents as a function of wages and the total months taken for both the pre and post-reform periods. We have also used lower numbers for this cost; these do not affect our conclusions.

Note that we have described only how parental leave choices affect income during the 17 month period. Future income, as discussed in Section 4.2, is also a function of these choices as described below.

6 Estimation

This section describes how we parameterize and estimate the model. Some model parameter values are directly taken from the data using model restrictions whereas the majority are estimated "internally" from the simulation of the model. There are a total of 18 internally estimated parameters which can be thought of as: 1) 2 parameters that govern preferences over current and future income; 2) 1 parameter that governs the curvature over the utility

[33] We also re-estimate the model using 0.8 and 0.9. Our results are very similar.

[34] We derive the figures for childcare costs as follows. First, the cost of providing public daycare is available from government reports (Skolverket (2010) and Skolverket (2000)). The annual costs amount to SEK 83,000 and SEK 117,500, respectively, or SEK 100,694 and SEK 122,696 in 2014 prices. Public daycare is, however, very subsidized - in 1999 parents paid 18% of the cost for the first child. The corresponding number in 2009 was 8%. This gives the monthly fee that parents pay (SEK 18,125 in 1999 and 9,815 in 2000), which is equal to 9.1% of the average wage of a non-university woman in the pre-period and 3.6% percent in the post period where the average wages of a non-university woman are from the wage sample (see Appendix Table A2).

We do not have data on the costs of private child care. Here we assume that the cost of private daycare equals the full (unsubsidized) cost public daycare. This would correspond to 50% of the average wage of a non-university woman in the pre-period and 45% in the post period. In our robustness checks, we have experimented with lower numbers for each: by taking 50% and 25% of the above figures and obtained similar results.

of time spent on parental leave, which differs by gender and education for a total of 4; and 3) 2 parameters associated with the evolution of the mean of the γ distribution and 1 associated with its variance, all of which can differ by gender and education for a total of 12. Table 2 summarizes the list of parameters.

6.1 Wage Parameters and Functional Forms

We assign monthly wages to an individual, based not only on that person's gender and education, but also on the education of their partner, as reflected in the couples data. In particular, we assume that, for each couple, the log wages one year prior to the birth of their first child are drawn from a bivariate normal distribution with parameters that differ by gender-household type and that are obtained directly from the data on log wages of each gender-household type in the couples sample.[35] The joint distribution for household wages takes the following form:

$$\begin{pmatrix} \log w_{m,t-1} \\ \log w_{f,t-1} \end{pmatrix} \sim N(\boldsymbol{\mu}, \Omega)$$

where

$$\boldsymbol{\mu} = \begin{pmatrix} \mu_w^{m,ij} \\ \mu_w^{f,ij} \end{pmatrix}, \quad \Omega = \begin{pmatrix} (\sigma_w^{m,ij})^2 & \rho^{ij}\sigma_w^{m,ij}\sigma_w^{f,ij} \\ \rho^{ij}\sigma_w^{m,ij}\sigma_w^{f,ij} & (\sigma_w^{f,ij})^2 \end{pmatrix}$$

and where ρ^{ij} is the correlation of mothers' and fathers' log wages. The parameters of this joint distribution are given in Appendix Table A3. When we calculate household income, all wages are normalized by the mean wages of non-university women in the pre period.[36]

Given each draw of $\log(w_{i,t-1})$, we construct future log wages using equation 3, the couple's parental-leave choices, and the wage-growth penalties in Table 1. That is, we calculate:

$$log(w_{i,t+3}) = \alpha_0^{g,e} + log(w_{i,t-1}) + \sum_{j=1}^{T^g} \beta_j^{g,e} \cdot \mathbf{1}[j = \text{plcat}_i] + \overline{\text{age}}^{g,e} + \bar{\phi}^{g,e} \tag{6}$$

where the year fixed effects have been replaced by the average of the coefficient values, $\bar{\phi}^{g,e}$, and the coefficients on the age dummies are replaced by the value of the coefficient on the

[35] These are monthly full-time equivalent wages.

[36] That is, after drawing log wages from the bivariate distribution, we first exponentiate each and then divide by the mean wage of non-university women in the pre-reform period.

median age dummy for their gender-education group, $\overline{\text{age}}^{g,e}$, since the model does not have several sample years nor heterogeneity in age. Appendix Table A4 reports the relevant values of $\alpha_0^{g,e}$, $\overline{\text{age}}^{g,e}$, and $\bar{\phi}^{g,e}$.

Lastly, we exponentiate the predicted future log wages, normalize as above, and sum both spouses' normalized wages to obtain total monthly household income $y_{ij}(t_m, t_f)$, net of their wage penalty as indicated by Table 1 given each parent's choice of parental leave time. It is worth noting that during this time wages grew substantially between $t-1$ and $t+3$: on average men's wage grew between 22% and 32% and women's wages between 18% and 24% in the pre period. In the post period, wages grew between 16% and 22% for men and by 13% and 17% for women.

We assume that utility over income (current and future) is given by CRRA preferences:

$$u(y) = \frac{y^{1-\eta}}{1-\eta} \tag{7}$$

where $\eta \geq 0$. The constant (B) multiplying the utility function over future household income in equation 4 is the same for all households.

Utility governing the time spent with a child during parental leave is given by:

$$h_{g,e}(t) = \frac{t^{1-\zeta_{g,e}}}{1-\zeta_{g,e}} \tag{8}$$

where the curvature is a function of both gender and education.

6.2 Internally Estimated Parameters and Estimation Procedure

The remaining 18 parameters are estimated internally using the method of simulated moments to match key empirical moments for both the pre- and post-reform periods. These moments are calculated using data with sample restrictions as described in Section 3 and are reported in Table 4 where we also give their model estimated values. These consist of, for each period and household type, the mean months of parental leave taken by men and women (2x8 moments), the standard deviation of these months (2x8 moments), the percentage of men who took zero months (2x4 moments), the mean household share of leave taken by men (2x4 moments), and the correlation, by household type, between the

Table 2: Description of Parameters

Categories	Parameters	Symbol	Value
External parameters	Mean of log wage draws by gender & hh type	$\mu_w^{g,ij}$	see Table A3
	SD of log wage draws by gender & hh type	$\sigma_w^{g,ij}$	
	Correlation coefficient of spousal log wage draws by hh type	ρ^{ij}	
	Wage growth (zero penalty) by gender & education		see Table A4
	Monthly wage replacement rate for paid parental leave	κ	0.85
	Monthly cost of childcare under age of 1	C_r	
	Pre-reform period		0.50
	Post-reform period		0.45
	Monthly cost of childcare over age of 1	C_u	
	Pre-reform period		0.091
	Post-reform period		0.036
Utility	Weight on utility over future average hh income	B	
	Curvature of utility function	η	
	Curvature of utility of time with children by gender & education	ζ_g^e	
Social norms	Intercept for mean of γ distribution by gender & education	$a^{g,e}$	
	Sensitivity of mean of γ distribution to peers by gender & education	$b^{g,e}$	
	SD of γ distribution by gender & education	$\sigma_\gamma^{g,e}$	

Note: Monthly costs of childcare are expressed as a percentage of the mean wage of non-university women. See section 5.2 for more details. The γ distribution is assumed to be normal. See text for definitions of all variables.

male to female household wage ratio and the within-household share of leave taken by men (2x4 moments), totaling 56 moments.[37] In particular, it is important for the model to capture the fact that, despite incentives to do so, a substantial share of men do not take any parental leave, which is why we choose to match the percentage of men who take zero months of parental leave.[38]

An overview of our estimation procedure is as follows. For each guess of the parameter vector, we compute model-generated moments by simulating households and solving for the steady-state distributions of parental leave taken for both pre- and post-reform periods. In the model simulation, the choice set for households consists of a discrete number of parental leave months for each spouse ($t_m, t_f \in \{0, 1, \ldots, 17\}$), subject to the constraint that the total leave taken per household does not exceed 17 months (with highly-paid leave only available for 12 months in the pre period and 13 in the post period, followed by at most 3 months

[37] The mean share is computed by first calculating the within-couple share of parental leave taken by the father and averaging these shares.

[38] Note that this is not an approximation as we report the proportion of men who did not apply for any parental leave.

of very low paid leave which we set to zero). Using the specification of household utility of Equation 4, the model identifies the combination of leave taken by men and women that maximizes household utility.

We describe the details of the estimation procedure in Appendix Section 10.4. The estimated parameters minimize the weighted sum of squared differences between model-generated and empirical moments, defined as:

$$\hat{\theta} = argmin_{\theta} \sum_{i=1}^{56} \omega_i \cdot \left(\frac{m_i^s(\theta) - m_i^d}{m_i^d} \right)^2$$

Here, m^d represents the 56 empirical moments, while $m^s(\theta)$ denotes the moments simulated under parameter vector θ. ω_i represents the weight assigned to moment i, with all moments receiving a weight of one except for standard deviations, which are given a weight of 0.5.

Table 3 reports the values of the estimated parameters along with their standard errors.[39]

Table 3: Estimated Parameters

Parameters	Symbol	Estimate			
Utility over future average hh income	B	2.652			
		(0.038)			
Curvature of utility function	η	1.099			
		(0.013)			
		Men		Women	
		n	u	n	u
Intercept for mean of γ distribution by gender & education	$a^{g,ij}$	-0.003	-0.004	0.043	0.093
		(0.000)	(0.000)	(0.001)	(0.001)
Sensitivity of mean of γ distribution by gender & education	$b^{g,ij}$	0.024	0.028	0.010	0.005
		(0.000)	(0.000)	(0.000)	(0.000)
SD of γ distribution by gender & education	$\sigma_{\gamma}^{g,ij}$	0.071	0.090	0.061	0.053
		(0.001)	(0.001)	(0.001)	(0.001)
Curvature of utility of time with children by gender & education	$\zeta^{g,e}$	0.865	0.843	0.602	0.585
		(0.009)	(0.008)	(0.006)	(0.006)

Note: See Table 2 for descriptions of all variables. Parameters $a^{g,ij}$, $b^{g,ij}$, $\gamma_{\sigma}^{g,ij}$, and $\zeta^{g,e}$ are estimated by gender and education. n denotes non-university and u university. The standard errors of the parameters are estimated using the method of empirical bootstrap.

[39]The standard errors are obtained by drawing new samples (with replacement) of the data, keeping the sample size of each household type as in the data. The parameters are then re-estimated to match the same moments as before. We do this 2000 times and then calculate the standard errors.

6.3 Estimation Results and Parameter Interpretation

As can be seen in Table 4, the model does a very good job matching the data.[40] The estimates for the mean months taken by men and women by household type are generally very close to the data moments as are the mean shares of parental leave taken by men.[41] The correlations produced by the model are all a bit smaller in absolute value than those in the data, suggesting that the variation in γ needed to match moments such as the percentage of men taking zero leave decreases the correlation between the household wage ratio and the share of leave taken by men.

Table 4: Estimation Targets: Model vs Data

Panel A: Pre-Reform Period

	Men				Women					
	Mean and SD Months		% Taking 0 Month		Mean and SD Months		Mean Share Taken by Men		Corr between Wage Ratio and Men's Share	
HH Type	Data	Model	Data	Model	Data	Model	Data	Model	Data	Model
nn	1.99	2.05	26.68	27.28	13.67	13.48	12.70	13.35	-0.09	-0.07
	(2.31)	(2.03)			(3.17)	(2.80)				
nu	2.62	2.54	21.40	22.32	12.79	12.53	17.29	16.85	-0.13	-0.09
	(2.37)	(2.25)			(3.36)	(2.82)				
un	2.31	2.30	28.18	25.61	13.27	13.53	14.73	14.58	-0.12	-0.08
	(2.49)	(2.24)			(3.41)	(2.78)				
uu	3.25	3.14	18.49	17.49	12.02	12.26	21.21	20.18	-0.13	-0.10
	(2.62)	(2.65)			(3.42)	(2.93)				

Panel B: Post-Reform Period

	Men				Women					
	Mean and SD Months		% Taking 0 Month		Mean and SD Months		Mean Share Taken by Men		Corr between Wage Ratio and Men's Share	
HH Type	Data	Model	Data	Model	Data	Model	Data	Model	Data	Model
nn	2.97	2.99	18.74	17.86	12.88	12.77	18.51	19.08	-0.10	-0.06
	(2.62)	(2.31)			(3.10)	(2.83)				
nu	3.83	3.75	12.15	11.50	12.16	11.85	23.95	24.07	-0.10	-0.10
	(2.70)	(2.43)			(3.26)	(2.84)				
un	3.45	3.51	15.28	15.46	12.34	12.37	22.03	22.14	-0.06	-0.08
	(2.64)	(2.62)			(3.35)	(2.96)				
uu	4.54	4.61	8.29	8.77	11.30	11.26	28.86	28.90	-0.11	-0.10
	(2.65)	(2.83)			(3.26)	(2.96)				

Note:. Standard deviation moments are in parentheses. Household types are denoted by the education level of the man first and the woman second, where n represents non-university and u represents university education.

[40] The weighted sum of squared errors is 0.70.

[41] Note that being close for means does not imply being close in shares as the mean shares taken by men are calculated couple by couple before being averaged.

It is useful to provide an interpretation of some of the parameters reported in Table 3. First note that η – the parameter governing the curvature of the utility function over income, common to all – is slightly greater than one and that the parameter weighting the utility of income three years later (B) is 2.65, indicating that household income three years later, which is a proxy for the household's entire future income stream, gets significantly more than twice the weight of household income during the parental leave period.[42] ζ – the parameter governing the curvature of the utility function over the amount of time spent on parental leave – is below one for all gender-education pairs but it is higher for men than for women indicating that, for the same γ, men's marginal utility from an additional month spent on parental leave is lower. Lastly, the b parameter governing the sensitivity of the mean of the γ distribution to changes in the mean parental leave of one's peers, is higher for men than for women, independent of education. This implies that, ceteris paribus, men will respond more to a change in their peer's parental leave behavior than women. Note that the standard deviation of γ is similar across education groups within gender, but greater for men than for women. The model requires this to accommodate the fact that the standard deviation of men's PL months is large relative to its mean in the data, whereas the standard deviation of women's PL months is small relative to its mean (see Table 4).

To understand how different households value parental leave, we need to evaluate the product of γ and $h(t)$ at some given t. We perform the following exercise: we calculate the percentage of household income an individual would be willing to sacrifice in order to obtain an additional month of leave assuming that γ takes the mean value for that gender and household type, $\mu_\gamma^{g,ij}$, and that they are endowed with the mean income for their household type, y^{ij}.[43] We arbitrarily choose to evaluate this at 6 months of parental leave which is high for men and low for women. That is, we solve for x below for each gender and household type:[44]

$$U(y^{ij}) - U(y^{ij}(1-x)) = \mu_\gamma^{g,ij} \cdot (h_{g,e}(7) - h_{g,e}(6))$$

[42] If we convert this into a constant stream of income for an additional 33 years as of year t+3, this implies a discount factor of 0.80.

[43] Household income here abstracts from any cost of parental leave – both from only receiving κ of one's monthly wage and from any childcare costs.

[44] Note that x can be solved for algebraically.

using the appropriate means of household incomes and $\mu_\gamma^{g,ij}$ which differ in the pre-reform and (endogenously) in the post-reform period as shown in the first row of each panel of Table 5.

Not surprisingly, as can be seen in the second row of Table 5, men of all household types are willing to sacrifice a much smaller percentage of household income than women in both periods. This percentage increases for men in the post period whereas it decreases for women. Given that mean household income increased for all types, the income effect increase couples' desires to increase parental leave. The fact that women's leave decreases must therefore be a response to other forces and forms part of the analysis in the next section.

It is also interesting to note that non-university men have a lower mean γ than university men when partnered with a woman of a given education, as can be seen from the first row of Table 5. This feature helps match the fact that non-university men take less parental leave than their university counterparts, given their partner's education. Similarly, in both periods, university women have a lower mean γ than non-university women, helping explain the fact that university women take less leave on average.

Table 5: Parameter Interpretation

Pre-Reform Period

	Men				Women			
	nn	nu	un	uu	nn	nu	un	uu
$\mu_\gamma^{g,ij}$	0.047	0.058	0.060	0.083	0.175	0.158	0.175	0.156
Additional Month at 6 Months (%)	0.97	1.30	1.39	2.06	6.23	5.96	6.39	6.00

Post-Reform Period

	Men				Women			
	nn	nu	un	uu	nn	nu	un	uu
$\mu_\gamma^{g,ij}$	0.069	0.087	0.093	0.124	0.168	0.154	0.164	0.151
Additional Month at 6 Months (%)	1.47	1.94	2.11	2.94	6.15	5.96	6.21	5.95

Note: The second row in each panel reports the percentage of average household income (where the latter varies by household type and pre vs post reform period) an individual of a given household type and gender would be willing to sacrifice in order to increase its parental leave from 6 months to 7. Household types are denoted by the education level of the man first and the woman second, where n represents non-university and u represents university education.

6.4 Identification and Sensitivity

We can gain insight into identification by performing two related exercises: (i) examine how changes in various parameters affect key equilibrium outcomes and, (ii) ask which moment

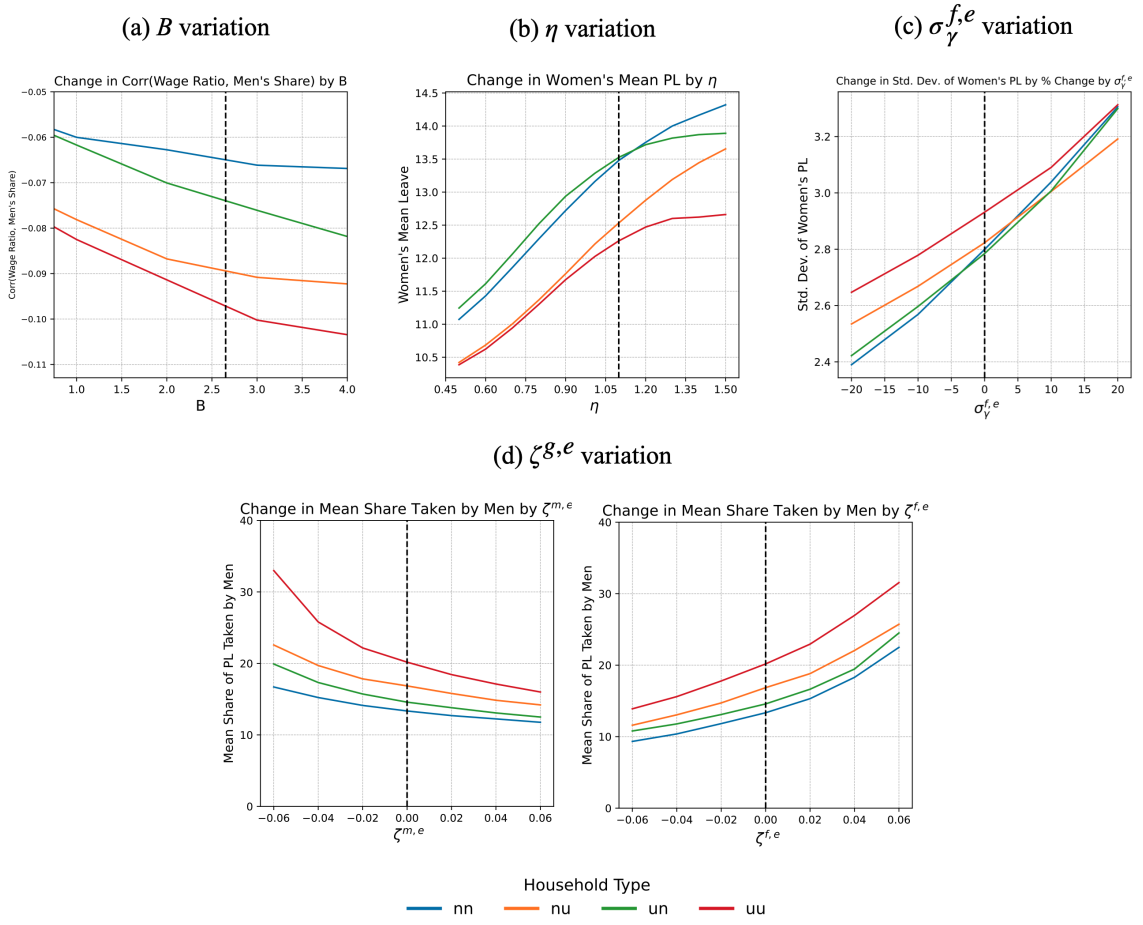
outcomes are most affected by a small change in a parameter value as in Chiplunkar and Goldberg (2024).

Figure 4 illustrates how large variation in $B, \eta, \sigma_\gamma^{f,e}$, and $\zeta^{g,e}$ affect given equilibrium outcomes in the pre-reform period, keeping all other parameters fixed. The dashed line indicates the estimated value of the parameter. As can be seen in Figure 4(a), higher values of B are associated with the correlation between the household wage ratio and the men's share of leave falling for all household types, i.e. increasing in absolute value. The intuition for this follows from higher values of B increasing the importance of *future* income and the opportunity cost of parental leave, thereby strengthening the conventional economic relationship between relative household wages and who foregoes income to stay home with a baby given that men tend to earn more and face higher wage penalties. Turning next to increases in η, as shown in Figure 4(b) this increases the mean leave of women. Intuitively, a higher η decreases the marginal utility of consumption (income), allowing the household to place relatively greater weight on time spent with the child and therefore increasing women's parental leave time. Next (c) graphs the effect of changes in the variance of the γ for women. As this varies by education, the x axis measures the percent deviation from its estimated value. Greater variances of the γ distributions not surprisingly increases the standard deviation of women's parental leave as since tastes for the latter as now more dispersed. Lastly, (d) shows the effect of changes in ζ – the curvature on the utility of time spent on parental leave – on men's mean share of parental leave. The figure on the left is for changes in $\zeta^{m,e}$ whereas the right is for $\zeta^{f,e}$. These parameters also vary by education so the x axis measures changes in the value of the parameter relative to its estimated value.[45] A higher ζ implies a lower marginal utility from parental leave. Hence, increasing $\zeta^{m,e}$ decreases men's share of total parental leave, whereas an increase in $\zeta^{f,e}$ increases their share as women want to spend less time on leave.

We can shed further light on identification by conducting an analysis similar to Chiplunkar and Goldberg (2024). We perturb each estimated parameter individually by 1%, holding all other parameters constant, and then measure the resulting absolute change in each simulated moment. This approach, akin to analyzing the Jacobian, highlights how sensitive

[45] These are not percentage changes since we need to avoid $\zeta > 1$ as a significant proportion of men take zero leave.

Figure 4: Parameter Variation



Note: The dotted black lines indicate estimated parameter values. For B and η, the x-axis represents the parameter value. In (c), we vary women's $\sigma_\gamma^{f,e}$, adjusting both education types simultaneously as percent changes from their estimated values. In (d), we vary men's $\zeta^{m,e}$ for both education types on the left panel and women's $\zeta^{f,e}$ for both education types in the right panel. In this case, however, the x axis shows the deviation from the estimate, with $x = 0$ marking the estimated value. Household types are denoted by the education level of the man first and the woman second, where n represents non-university and u represents university.

the moments are to each parameter and, consequently, how these parameters are identified.

As shown in Appendix Figure 6, parameters related to income, specifically B (weight on utility from future income) and η (curvature of the utility function over income), primarily affect income-related moments, such as the correlation between the household male-to-female wage ratio and the man's share of leave. This underscores how these parameters are primarily identified through economic relationships tied directly to leave decisions. The

preference parameters σ_γ, which capture variance in the preference for spending time with children, naturally affect the standard deviation of parental leave durations for both men and women. They also influence correlation moments, because adding preference-based heterogeneity makes leave choices less directly tied to income, further weakening observed income-leave correlations. Consequently, σ_γ is identified by the extent of non-income-driven heterogeneity in parental leave behavior. We can also examine this table using its rows to ask which parameters most influence particular moments. In this light, the ζ parameters, which capture the curvature of utility derived from parental leave months, have the greatest quantitative effects on the percentage of men taking zero leave and the mean leaves of both men and women. This makes sense as the taste distribution for leave implies that a significant fraction of men do not wish to take leave and the curvature of the utility with children will then have a large effect on the proportion that chooses a corner solution of zero as well as mean leaves.

7 Economic Incentives and Cultural Change

We can now turn to one of the principal concerns of this paper: determining how different changes in the environment – regulations, income-related changes as well as endogenously changing social norms – contributed to the evolution of parental leave take-up from the pre-reform distribution to the post-reform one.

We do this by first introducing the change in the parental leave regulation on its own, keeping all other economic factors and preferences unchanged (i.e, keeping the $\mu_\gamma^{g,ij}$ at their original pre-reform levels). Next, we examine the effect of the reform along with all income related changes, still keeping the $\mu_\gamma^{g,ij}$ at their original pre-reform levels. After that we drop the income-related changes and, along with the reform, allow the $\mu_\gamma^{g,ij}$ to respond to the reform. Finally, we allow all changes. In all cases we solve for the steady state given the restrictions of each decomposition.

Table 6: From Pre to Post Reform

Moments	Pre-Reform Simulated Moments	Reform Only	Income	Social Norms	Post-Reform Simulated Moments
			nn		
Mean months of men	2.05	2.25	2.33	2.64	2.99
SD months of men	2.03	2.01	2.11	2.15	2.31
Mean months of women	13.48	13.42	13.40	13.04	12.77
SD months of women	2.80	2.63	2.71	2.71	2.83
% Men taking 0 month	27.28	27.26	27.06	21.39	17.86
Mean share taken by men	13.35	14.41	14.81	16.91	19.08
Corr between wage ratio & men's share	-0.07	-0.07	-0.06	-0.07	-0.06
			nu		
Mean months of men	2.54	2.74	2.84	3.17	3.75
SD months of men	2.25	2.22	2.30	2.31	2.43
Mean months of women	12.53	12.59	12.56	12.28	11.85
SD months of women	2.82	2.69	2.82	2.73	2.84
% Men taking 0 month	22.32	22.34	22.05	16.07	11.50
Mean share taken by men	16.85	17.84	18.42	20.50	24.07
Corr between wage ratio & men's share	-0.09	-0.10	-0.09	-0.10	-0.10
			un		
Mean months of men	2.30	2.44	2.58	2.70	3.51
SD months of men	2.24	2.21	2.30	2.32	2.62
Mean months of women	13.53	13.50	13.33	13.25	12.37
SD months of women	2.78	2.64	2.75	2.71	2.96
% Men taking 0 month	25.61	25.62	25.88	21.84	15.46
Mean share taken by men	14.58	15.28	16.21	16.90	22.14
Corr between wage ratio & men's share	-0.08	-0.08	-0.08	-0.08	-0.08
			uu		
Mean months of men	3.14	3.27	3.44	3.62	4.61
SD months of men	2.65	2.60	2.62	2.70	2.83
Mean months of women	12.26	12.33	12.22	12.07	11.26
SD months of women	2.93	2.81	2.91	2.84	2.96
% Men taking 0 month	17.49	17.56	17.69	14.08	8.77
Mean share taken by men	20.18	20.74	21.81	22.84	28.90
Corr between wage ratio & men's share	-0.10	-0.10	-0.09	-0.11	-0.10

Note: Column 2 introduces the reform but no other changes. Column 3 allows, in addition to the reform, all income related variables (wage parameters, wage penalties, and childcare costs) to take their post-period values as given in Tables 1 and A4. Social norms (the $\mu_\gamma^{g,ij}$) are kept constant at the pre-reform values. Column 4 returns to the income-related variable values of the pre-reform period but now allows the values of the $\mu_\gamma^{g,ij}$ to evolve endogenously. Column 5 incorporates all changes. Household types are denoted by the education level of the man first and the woman second, where n represents non-university and u represents university education.

The Effects of the Reform Only

Recall that the reform increased the number of daddy months from 1 to 2 and the total earnings-related months of parental leave from 12 to 13. Consequently, it did not affect the maximum number of months a woman could take with full parental leave pay – they were still 11–while it increased the number of months a man would "throw away" by not taking parental leave. The consequences of only the reform, keeping all income-related variables at their pre-reform levels and not allowing culture to change are shown in column 2 of Table 6, titled "Reform Only."

For men, the reform on its own accounts for a non-negligible portion of the total increase in their parental leave spells of around one-fifth of a month. It generates from 8 to 11% (for uu men) to 21% (for nn men) of the total increase in their mean parental leave (i.e, of the increase from the pre-reform to the post-reform simulated moments). For women, the effect is varied, with university women slightly increasing and non-university women slightly decreasing their leaves. These results can be explained by noting that the reform increased the number of paid parental leave months reserved (de facto) solely for men, which increased their incentive to take more time without changing women's incentives in a substantial manner.

The Reform and all Income-Related Changes

Next, we introduce all the income-related changes from the post period in addition to the reform, still keeping the $\mu_{\gamma}^{g,ij}$ at their original pre-reform levels. These variables are the wage parameters and wage penalties (see Appendix Tables A3 and A4) and childcare costs (see Table 2). Over this time period, wages grew substantially for both men and women. Childcare costs also changed, decreasing for public childcare and increasing for private. The results are shown in column 3 titled "Income."

The incremental effects from all income-related changes generates an additional increase in men's leave ranging from from 8 to 11% of the total increase in their mean parental leave (i.e, of the increase from the pre-reform to the post-reform simulated moments). For women, the effect is a bit more varied, decreasing women's average leave from 3 to 15%, but with a similar overall average as men of 8.25%. As shown in Appendix Table A6 which

examines how each income-related change separately affects parental leave, for men the most important dimension was the decrease in wage penalties whereas for women they all played a minor role.

It is important to note that, for both of these experiments, the percentage of men who are taking zero months of leave barely changes. These men have very low draws of γ and have almost no incentive to change their behavior in response to the reform and changes in wage parameters.

The Reform and Endogenous Cultural Change

Next, we examine the effect of the reform keeping the income-related parameters constant at their pre-reform levels, but allowing the mean of the γ distributions to change endogenously in response to the reform and solving for the new steady state. The effects of this are substantially greater than in the prior experiments. Allowing culture to change accounts for an additional increase over the effect of the reform that ranges from 22% (un) to 41% (nn) of the final increase. For women, the additional contribution is also very large, especially for women with a non-college partner as it accounts for 45 (nn) and 53% (nu) of their total leave decrease.

For the first time, there is a significant effect on the percentage of men who take zero leave. This is especially large for men without a college education. The implication is clear: these men change their behavior because the reform causes other men to take greater parental leave, thereby influencing the behavior of their peers.

It is reassuring that there is poll evidence regarding changes in attitudes. In 2002, 44% of university women, as opposed to 17% of non-university women, strongly disagreed with the statement "Family life suffers if the woman works full time." The corresponding figures for university men and non-university men were 26% and 14%, respectively. In 2012, 55% of university women and 33% of non-university women disagreed with the statement, as did 42% of the university men and 24% of the non-university men.[46]

[46]Source: simple averages calculated from ISSP (The International Social Survey Programme), 2002, 2012 . See ISSP Research https://doi.org/10.4232/1.12661).

Putting it all Together

The last column includes all changes in the environment and is therefore identical to the post-reform simulated moments. Including all income-related changes in addition to allowing social norms to change endogenously (column 5 vs 4), has significant quantitative consequences. It accounts for a further 37 (nn) to 67% (un and uu) increase in parental leave, much larger than the independent incremental effect on the reform from either income-related or social norms related changes. The additional effect on women's parental leave is even larger, ranging from 39 (nn) to 81% (uu) of the final decrease in their parental leave. that of women. Overall, the extra kick from allowing both income and social norms to change comes from the dynamic amplification of the otherwise static change in incentives from the former. Endogenously changing social norms magnifies the response to income-related changes beyond the response stemming from the reform itself.

8 Delving Deeper

We next turn to exploring two facets of our results. First, we examine how the model-derived magnitudes of the effects of the reform only on parental leave compare with those obtained via a regression discontinuity estimate. Second, we examine how our results would differ if a key parameter – the correlation of household wage ratios and men's share of leave – were radically different. Second,

The Short-Run Effect of the Reform

In the prior section we decomposed the consequences of the reform by keeping constant all income-related variables as well as preferences at their pre-reform levels and imposing the reform. This is what we called the effect of PL choices of only the reform. Subsequently we introduced the income-related changes and allowed social norms to evolve endogenously to understand their contributions. Here we return to examining the effect of only the reform by using the discontinuity it introduced and contrasting it with the model-derived implications. As we show below, the model does fairly well in replicating the short-term effect of the reform.

We exploit the discontinuity introduced by the 2002 reform to estimate its impact on parental leave uptake using a sharp regression discontinuity (RD) design. To be consistent with our model estimation, we measure parental leave as the number of leave months taken within the first 16 months following childbirth and use our couples sample but restricted to births in 2001 and 2002. The cutoff is defined at January 2002, and a monthly running variable is used. Our RD sample consists of 6,059 couples, with monthly observations ranging from approximately 200 to 300 couples. Due to the small sample size, we do not differentiate by education or by couple type, but estimate the discontinuity at the cutoff (January 2002) separately for fathers and mothers.

The RD equation is given by:

$$PL_i = \alpha + \delta D_i + f(X_i) + u_i$$

where PL_i is the number of parental leave months taken within the 16 months following childbirth. The indicator variable D_i equals 1 if the child was born in 2002 and 0 if the child was born in 2001. The running variable X_i ranges from -12 to 11, taking the value 0 at the cutoff (January 2002). It represents months relative to the reform date, with negative values for births in 2001 and positive values for births in 2002. Details are given in Appendix Section 10.7.

Appendix Figure 7 presents the RD plots for a polynomial fit of order 0 (mean) and 1 (linear). These plots show a discontinuity in fathers' parental leave uptake of 0.388 months (mean – $p = 0$ in the polynomial specification) or of 0.268 in the linear specification ($p = 1$) though the last is statistically significant at the 10% level. In contrast, for mothers there is no clear evidence of a discontinuity associated with the 2002 reform.[47] The RD estimates

[47] Avdic and Karimi (2018), while primarily focused on the effect of the first daddy month reform of 1995, also study the 2002 reform. They use, however, the first eight years after birth to measure parental-leave uptake. They find that the 2002 reform increased total parental leave over the children's first 8 years by roughly 25 days per child, about 15 of which were taken by fathers. Duvander and Johansson (2012) also study the 2002 reform and find that both mothers and fathers each increased their leave by an average of six to seven days, based on a 24-month follow-up after birth. Focusing on fathers, we find a coefficient of 0.388, which implies about 11 to 12 extra days of leave (0.388×30). Our estimate is smaller than that of Avdic and Karimi (2018), likely reflecting our shorter, 16-month post-birth observation window. On the other hand, our estimate is somewhat larger than the one presented in Duvander and Johansson (2012). This likely reflects the fact that we use first births rather than all births and that university fathers, who are somewhat overrepresented in our couples sample, are often matched with university mothers and are therefore likely to be more responsive to the incentives embedded in the 2002 reform.

are reported in Appendix Table A7.

We can use our estimated model to simulate mean parental-leave uptake for men and women in 2001 and 2002, paralleling our RD design. To replicate the pre-reform environment in 2001, we impose the observed 2001 log-wage distribution by gender and household type, holding all other primitives – within-household wage correlations, wage-penalty parameters, and childcare-cost parameters–at the pre-reform levels used to estimate our model. We then solve for the model's steady state and compute mean parental-leave uptake by gender–household type. This is the model based counterpart of the pre-reform enviroment in 2001.

Next, to capture the immediate impact of the 2002 reform, we introduce one additional month of paid leave for fathers and increase the highly paid months to 13, but keep all income-related inputs and preferences (cultuat their 2001 pre-reform values. This exercise is analogous to the "Reform Only" column in Table 6, but now using the 2001 log wage distribution by household type rather than the average for the period. Finally, because the model produces outcomes at the gender–household-type level but the RD estimates report simple gender-specific means, we aggregate the simulated results using period-specific household-type weights derived from the RD sample shares in 2001 and 2002. The sample sizes and resulting weights are reported in Table A8.

Table 7 reports the results of this exercise. As shown in the table, men witnessed an increase of 0.22 months whereas women saw a negligible decrease of 0.03 months of parental leave. These results are close to the estimates obtained in our RD, showing that our model is able to do a good job of replicating the immediate effects of the 2002 reform.

Table 7: RD Simulation Results

	Men	Women
Pre-Reform, 2001	2.43	13.00
Post-Reform, 2002	2.64	12.97
Difference	0.22	-0.03

Note: This table reports the model-simulated mean months of parental leave taken by fathers and mothers in 2001 (pre-reform) and the 2002 (post-reform). Simulations mirror the RD environment by imposing the observed 2001 log-wage distribution by gender and household type, implementing the 2002 policy change, and aggregating outcomes using the period-specific household-type weights from Table A8. The "Difference" row indicates the change in leave uptake (post minus pre).

A Counterfactual Correlation

A legitimate question is whether the model allows social norms to play a much smaller role and income-related changes to play a larger role. Our answer is clearly yes. Had the absolute level of the correlation between the household ratio of male to female wages and the men's share of leave been significantly greater, income-related changes would have played a dominant role. Appendix Table A5 reports the results of re-estimating the model using a correlation target of -0.60 for all households rather than its current value which ranges from -0.07 to -0.10 depending on household type.[48] In that counterfactual estimation, the reform has only a small effect on parental leave. The additional contribution of income-related changes is large, ranging from 49 to 89%. Furthermore, while the percentage of men taking zero leave changes very little for no-university men, it significantly decreases it for university men. By contrast, social norms' additional contribution to the reform is small and has very little effect on the percent of men who take zero leave. However, as can be seen by comparing columns 5 and 3, the dynamic effects of social norms remain important, contributing from 28 to 43% of the change in addition to the effect of the income-related changes. Thus, the model could allow a greater role for income and a smaller role for culture had certain key moments been different.

9 Increasing Gender Equality in Parental Leave

Achieving greater gender equality by having men "assume an equal share of caring responsibilities" is an explicit objective of the paternity and parental leave directives issued by the European Union.[49] Following the example of Sweden in earmarking parental leave months for each parent, the directive instructs its members to endow each worker with the right to four months of paid parental leave, of which 2 months are not transferable across parents.[50] In Sweden, during the period we study, this reform increased men's share of leave, but still

[48] This estimation has a significantly larger sum of squared errors of 1.72 as opposed to 0.70.

[49] The quote is from the EU directive. See the summary of this directive (EU Directive 2019/1158 on "Work-life balance for parents and carers"): https://eur-lex.europa.eu/legal-content/EN/LSU/?uri=CELEX:32019L1158.

[50] The directive does not limit the generosity of the leave system, permitting countries to provide greater lengths of paid leave.

left it below 30% and, for some household types, below 20%. In this section we make use of one of the main virtues of an estimated model by employing it to study how alternative policies could affect gender equality.

We evaluate all exercises assuming that the economic parameters remain the same as those under the two-daddy months policy, i.e., with the same wage distribution parameters, wage penalties, and childcare costs as in the post reform environment.[51] Preferences for spending time with the child are also at their steady state levels under the two-daddy-months policy.

For all our counterfactual exercises, we conduct two experiments. First we keep culture fixed (i.e., assume that the $\mu_\gamma^{g,ij}$ stay constant at the 2-daddy-months reform levels), and then we report the new steady states that result from allowing social norms to endogenously change. In each table, panel A reports the simulated moments by household type and panel B reports the $\mu_\gamma^{g,ij}$.

9.1 The Role of the Gender Wage Gap

On average, university women earn 84% of university men's wage, and non-university women earn 86% of their male counterparts a year prior to the birth of their first child.[52] Before studying counterfactual policies, we can examine the role of the gender wage gap in parental leave choices of men and women. To do so, we give women the same wage distribution as men in their household type, e.g. non-university women partnered with university men now have their wage distribution.[53] The correlation of wages and wage penalties by education and gender are kept unchanged from their values in the post-reform period.

Table 8 reports the results of this exercise. For ease of comparison, column 1 reproduces the results of the two-daddy month reform. Not surprisingly, the effects of the exercise is to increase men's parental leave and decrease that of women, but in the absence of cultural change the effects are relatively small (column 2), only changing men's share of parental leave by around 1 percentage point. Once preferences are allowed to change endogenously

[51] As we do not have a model of wage determination or signaling nor one of sorting in the "marriage market," the quantitative exercise assumes that these do not respond to changes in parental leave behavior.
[52] Calculated from the couples sample, year t-1 of the post-reform period.
[53] Recall that wage distributions do not simply vary by education, but also by household type.

(column 3), the effects are greater in magnitude but nonetheless the share of parental leave taken by men remains under 30% except for uu couples where it is 32.7%. The intuition for this result is that although income-related factors play a non-negligible role in driving parental leave, the low correlation between wage ratios and men's leave shares indicates that preferences trump income considerations. Thus even when women and men share a common wage distribution, this is an insufficiently important force for promoting gender equality in parental leave uptake.[54]

9.2 Counterfactual Policies

We start by considering a policy that lowers the cost of childcare by offering public childcare prior to the age of one at the same subidized price as after that age – over a twelvefold decrease. Next we consider an "endowment policy", which gives each parent 6 months of non-transferable paid leave, with an additional 1 month that can be taken by either parent. Both Iceland and Finland have parental leave regimes similar to this.[55] In Iceland, for children born as of 2021, each parent is endowed with 6 months of parental leave benefits of which 6 weeks can be transferred to the other parent. In Finland, for children born as of September 4, 2022, each parent is endowed with 160 working days of parental leave benefits of which 63 days can be transferred. Lastly, we examine a stricter "equal sharing" policy under which a parent can take up to six months of paid leave but the months in excess of those taken by their partner are forfeited. There is one paid month, however, that can be taken by either partner independently of whether it is matched.[56] In all cases, anyone can take the three months of low-paying leave (zero pay).

As can be seen in columns 4 and 5 of Table 8, providing inexpensive childcare during the child's first year has an insignificant effect on men's mean months of parental leave. Women's mean months only very slightly decrease and men's mean leave months remain virtually unchanged. Note that there is no room for social norms to change in a significant fashion

[54]Furthermore, the extensive part of preferences – the $h(t)$ functions – are invariant and imply that men have a significantly lower marginal utility of leave. One might also imagine that this would change over time, bringing men's and women's preferences closer together.

[55]See norden.org/en/info-norden/parental-benefits-iceland and norden.org/en/info-norden/parental-benefits-finland.

[56]The extra month in both this policy and the endowment policy keeps the maximum number of paid months at the same level as in the second daddy month policy.

given the lack of behavioral responses to this policy, which explains why both columns show basically the same results. Thus, according to our model the significantly higher cost of childcare in the child's first year is not responsible for the gender asymmetry in parental leave and changing its price would not result in greater sharing of care responsibilities.

Table 8: Wages and Childcare

Simulated Moments

Moments	Post-Reform Simulated Moments	Equal Wages Unchanged $\mu_\gamma^{g,ij}$	Equal Wages Steady State	Reduction in Childcare Unchanged $\mu_\gamma^{g,ij}$	Reduction in Childcare Steady State
			nn		
Mean months of men	2.99	3.08	3.40	2.99	2.99
SD months of men	2.31	2.35	2.44	2.30	2.30
Mean months of women	12.77	12.52	12.19	12.74	12.73
SD months of women	2.83	2.79	2.87	2.85	2.86
% Men taking 0 month	17.86	17.69	14.66	17.93	17.97
Mean share taken by men	19.08	19.72	21.85	19.08	19.10
Corr wage ratio & men's share	-0.07	-0.09	-0.08	-0.06	-0.06
			nu		
Mean months of men	3.75	3.81	4.02	3.74	3.74
SD months of men	2.43	2.46	2.48	2.43	2.43
Mean months of women	11.85	11.72	11.55	11.82	11.81
SD months of women	2.84	2.80	2.81	2.86	2.86
% Men taking 0 month	11.50	11.44	9.93	11.56	11.60
Mean share taken by men	24.07	24.52	25.81	24.07	24.08
Corr wage ratio & men's share	-0.10	-0.11	-0.11	-0.10	-0.10
			un		
Mean months of men	3.51	3.70	4.58	3.51	3.52
SD months of men	2.62	2.73	2.97	2.61	2.62
Mean months of women	12.37	12.00	11.17	12.33	12.31
SD months of women	2.96	2.92	3.12	3.00	3.00
% Men taking 0 month	15.46	15.32	9.47	15.49	15.46
Mean share taken by men	22.14	23.39	28.98	22.20	22.27
Corr wage ratio & men's share	-0.08	-0.10	-0.10	-0.07	-0.07
			uu		
Mean months of men	4.61	4.76	5.23	4.62	4.63
SD months of men	2.83	2.88	2.93	2.83	2.83
Mean months of women	11.26	11.02	10.64	11.23	11.22
SD months of women	2.96	2.89	2.91	2.97	2.97
% Men taking 0 month	8.77	8.73	6.46	8.78	8.74
Mean share taken by men	28.90	29.82	32.68	28.97	29.05
Corr wage ratio & men's share	-0.10	-0.11	-0.11	-0.10	-0.10

Note: Columns 2 and 3 report the moments obtained by equalizing women's log wage distribution to men's, within a household type. Columns 4 and 5 report the moments from reducing monthly childcare costs in the first 12 months from 45 percent of the average monthly wage of a non-college woman in post-period to 3.6 percent. Household types are denoted by the education level of the man first and the woman second, where n represents non-university and u represents university education.

Next, we examine counterfactual policies that change the number of months of non-transferable parental leave and the rules governing these. The consequences of these policies are reported in Table 9.

Table 9: Alternative Policies

Panel A: Simulated Moments

Moments	Post-Reform Simulated Moments	Endowment Unchanged $\mu_\gamma^{g,ij}$	Endowment Steady State	Equal Sharing Unchanged $\mu_\gamma^{g,ij}$	Equal Sharing Steady State
			nn		
Mean months of men	2.99	4.40	6.47	5.28	6.70
SD months of men	2.31	2.46	1.68	2.09	1.37
Mean months of women	12.77	11.50	9.42	10.67	9.26
SD months of women	2.83	2.65	2.40	2.02	2.05
% Men taking 0 month	17.86	17.64	1.74	12.82	0.81
Mean share taken by men	19.08	27.44	41.29	32.58	42.38
Corr wage ratio & men's share	-0.07	-0.13	-0.16	-0.11	-0.19
			nu		
Mean months of men	3.75	5.16	6.64	5.77	6.83
SD months of men	2.43	2.24	1.58	1.75	1.33
Mean months of women	11.85	10.57	9.46	10.03	9.32
SD months of women	2.84	2.35	1.92	1.72	1.67
% Men taking 0 month	11.50	11.28	1.38	7.45	0.66
Mean share taken by men	24.07	32.55	41.41	36.14	42.43
Corr wage ratio & men's share	-0.10	-0.14	-0.21	-0.13	-0.22
			un		
Mean months of men	3.51	4.53	6.87	5.41	7.24
SD months of men	2.62	2.56	2.07	2.23	1.66
Mean months of women	12.37	11.39	9.20	10.65	8.97
SD months of women	2.96	2.73	2.51	2.21	2.04
% Men taking 0 month	15.46	15.29	1.93	11.96	1.03
Mean share taken by men	22.14	28.26	43.17	33.23	44.89
Corr wage ratio & men's share	-0.08	-0.16	-0.20	-0.17	-0.24
			uu		
Mean months of men	4.61	5.57	7.01	6.16	7.36
SD months of men	2.83	2.39	1.98	1.94	1.59
Mean months of women	11.26	10.36	9.29	9.87	9.05
SD months of women	2.96	2.38	2.06	1.83	1.73
% Men taking 0 month	8.77	8.66	1.79	6.50	0.89
Mean share taken by men	28.90	34.64	43.02	38.03	44.88
Corr wage ratio & men's share	-0.10	-0.17	-0.23	-0.17	-0.26

Panel B: Mean values of $\mu_\gamma^{g,ij}$

	Post-Reform Simulated Moments	Endowment Unchanged $\mu_\gamma^{g,ij}$	Endowment Steady State	Equal Sharing Unchanged $\mu_\gamma^{g,ij}$	Equal Sharing Steady State
Men					
nn	0.07	0.07	0.15	0.07	0.16
nu	0.09	0.09	0.16	0.09	0.16
un	0.09	0.09	0.19	0.09	0.20
uu	0.12	0.12	0.19	0.12	0.20
Women					
nn	0.17	0.17	0.13	0.17	0.13
nu	0.15	0.15	0.14	0.15	0.14
un	0.16	0.16	0.13	0.16	0.13
uu	0.15	0.15	0.14	0.15	0.14

Note: Columns 2 and 3 study an equal endowment policy of 6 months for each parent (with an additional month allocated freely) whereas columns 4 and 5 study an equal share policy that only pays parental leave for the minimum months taken by a parent, with one month not subject to that rule. The first column for each policy keeps the $\mu_\gamma^{g,ij}$ at the steady-state values of the 2-daddy months policies. The second column for each policy allows preferences to adjust endogenously. Household types are denoted by the education level of the man first and the woman second, where n represents non-university and u represents university education.

Columns 2 and 4 report the moments resulting from the endowment and equal-sharing policies, respectively, assuming that cultural preferences are unchanged. Even with no change in preferences, these policies increase men's average parental leave substantially. Under the endowment policy, men's mean parental leave increases from around 1 to 1.5 months. The fall in women's parental leave is similarly large. The equal sharing policy produces even larger static effects for both men and women. These changes work in the same direction to substantially increase men's share of parental leave under both policies but by substantially more under equal sharing.

It is easy to understand why the equal sharing policy produces greater changes in men's and women's parental leaves than the endowment policy. When a father with, say, 3 months of parental leave increases it by one month under the endowment policy, this does not have any repercussions for the mother, except via the adjustment of household income.[57] When a similar father increases his parental leave by one month under the equal sharing policy, it not only has the same repercussions for household income as with the endowment policy, but in addition allows the mother to take an additional month of paid leave. If her parental leave had been more than a month longer than the father's leave, this would increase household income by an additional $.85w_f$ as now this month is compensated. Consequently, men have an additional incentive to take parental leave under the equal sharing policy that is absent under the endowment policy.[58] A similar logic explains why women decrease their parental leave more under the equal sharing versus the endowment policy. The difference in incentives created by these two policies is most clearly seen in the share of men who take zero months of parental leave. This share barely changes with the endowment policy, but decreases by several percentage points under the equal share policy as, by not taking any leave, these men are allowing their partners to take only one month of paid leave.

Lastly, we compare the steady-state implications of these two policies allowing preferences to change endogenously, generating the moments in columns 3 and 5. The final results now look very similar across both policies. On average, under both policies non-university men end up taking around 6.7 months of parental leave whereas university men take around 7.1

[57] Household income decreases by $0.15w_m$ as only 85% of his wage is paid during leave, but it also increases by the extra month of childcare expenditures saved by the couple.

[58] If both parents were taking more than 6 months of parental leave, the two policies would have identical incentives as there are no additional months of paid leave available.

months. The percentage of men who take zero months of leave is under 2% for all but men. Under both policies, women take similar amounts of leave. Thus, although incentives with invariant preferences gave rise to quantitatively significant differences in behavior under the two policies, evolving social preferences imply that the final results are very similar. Finally, while neither policy obtains a 50-50 split of parental leave, they all surpass 40%, with non-university men averaging 42% and university men averaging 44%.

Panel B of Table 9 shows the mean values of the $\mu_\gamma^{g,ij}$ under the four policies: the post-period (i.e., 2 daddy months policy, which is the same in columns 2 and 4), and those in the steady-states of the different policies (columns 3 and 5). The steady state values under the endowment and equal share policies are markedly different from those of the 2-daddy-months steady state, especially for men. It is interesting to note that men's $\mu_\gamma^{g,ij}$s are now higher than women's. This does not imply, however, that men prefer to spend more time with children than women. The curvature of the utility function over time with one's child also plays an important role. As can be seen in table 3, men have a substantially higher value of ζ than women, independently of education, implying that their marginal utility of spending time with children is lower.

10 Conclusion

Even in countries in which gender equality is more advanced, care responsibilities continue to follow traditional gender roles. This paper studies the 2002 parental leave reform in Sweden, which introduced a second reserved month for each parent (a second daddy month). This reform increased the share of parental leave months taken by men, but the division remained far from equal and, even today, men's share does not exceed 30%. What reforms might work to change this? To answer this question and to quantify the importance of intermediating mechanisms, we develop a unitary household model in which individuals care about consumption (both during the parental leave period and in the future) and also derive utility from spending parental leave time with their child. The latter is not a primitive, however, as it is influenced by the behavior of one's peers, making it an equilibrium object.

We distinguish individuals by their education (university and non-university), implying

four types of households. Taking wage parameters including estimated wage penalties associated with different lengths of parental leave as given, we estimate the model to match key parental leave moments both pre- and post-reform for each household type. We then study the quantitative importance of different mechanisms for the post-reform outcomes. We first examine the incentives introduced by the reform itself keeping all else constant. Next, we introduce the income-related changes along with the reform (keeping social norms constant). Lastly, we allow preferences to change both in response to the reform only and to both the reform and income-related changes. We find that the reform alone and income-related changes in addition to the reform played a significant but quantitatively relatively modest role in the absence of endogenous cultural change. The evolution of preferences both in response to the reform and in response to income-related changes played the largest role for both men and women from all household types.

Lastly, we use the estimated model to evaluate the role of the gender wage gap and three alternative reforms: providing low-cost childcare before the age of one, giving each parent 6 months of non-transferable parental leave (the "endowment" policy), and only paying for the months of parental leave that do not exceed those taken by one's partner (the "equal share" policy). We find that equating wage distributions of men and women by household type affects parental leave, but still leaves men's share of parental leave below 30% for all household types except uu couples where it increases to 32.7%. Decreasing the cost of childcare has almost no effect on men's share of parental leave indicating that it is unlikely that such a policy would produce significant changes. Both the endowment and the equal share policies, on the other hand, have similar steady-state consequences, increasing men's share significantly. Although it is possible to argue that how partners share parental leave time is of second-order concern, greater sharing of parental leave responsibility is likely to have significant spillover effects. As fathers take more time caring for children over the life-cycle, their demand for more flexible jobs should increase. This in turn could decrease the "motherhood penalty" in earnings that is found virtually everywhere. Furthermore, as shown by Kotsadam and Finseraas (2013) and Farré et al. (2023), it may have important intergenerational consequences that further change attitudes and economic outcomes.

Appendix

10.1 Summary Statistics

Table A1: Distribution of Household Types

	nn	un	nu	uu	Total
Pre Period	5,266	880	1,757	2,023	9,926
	(53.1%)	(8.9%)	(17.7%)	(20.4%)	
Post Period	4,200	975	3,326	5,295	13,796
	(30.4%)	(7.1%)	(24.1%)	(38.4%)	

Note: Tabulation of the couples samples. Household types refer to education of man first, woman second, n = non-university, u = university. See text for exact definition of education and sample.

Table A1 gives the distribution of household types in the couples samples by period.

Table A2 presents summary statistics for the wage sample and couples sample discussed in Section 3 above.

Table A2: Summary Statistics

	Wage Sample		Couples Sample	
	Pre-Period	Post-Period	Pre-Period	Post-Period
	Men - Non-University			
Monthly Wage $(t+3)$	24704	31361	24568	31212
Monthly Wage $(t-1)$	19758	26801	19644	26663
PL (months)	2.15	3.26	2.15	3.35
Age	31.2	31.7	31.4	32.1
Obs	16511	17450	7023	7526
	Men - University			
Monthly Wage $(t+3)$	33404	40502	32721	39966
Monthly Wage $(t-1)$	24121	32544	23892	32322
PL (months)	2.84	4.27	2.96	4.37
Age	33.3	33.3	33.2	33.4
Obs	6069	12961	2903	6270
	Women - Non-University			
Monthly Wage $(t+3)$	19830	25734	20058	26279
Monthly Wage $(t-1)$	16606	22622	16808	22987
PL (months)	13.9	13.1	13.6	12.8
Age	29.2	29.0	29.4	29.3
Obs	15515	13491	6146	5175
	Women - University			
Monthly Wage $(t+3)$	24933	32054	25039	32258
Monthly Wage $(t-1)$	19612	27096	19712	27132
PL (months)	12.7	12.0	12.4	11.6
Age	30.7	31.4	30.7	31.5
Obs	7902	19212	3780	8621

Note: Monthly full-time equivalent wages are expressed in 2014 SEK. See the text for definition of PL months. Source: Monthly full-time equivalent wages are taken from the Wage Structure Statistics.

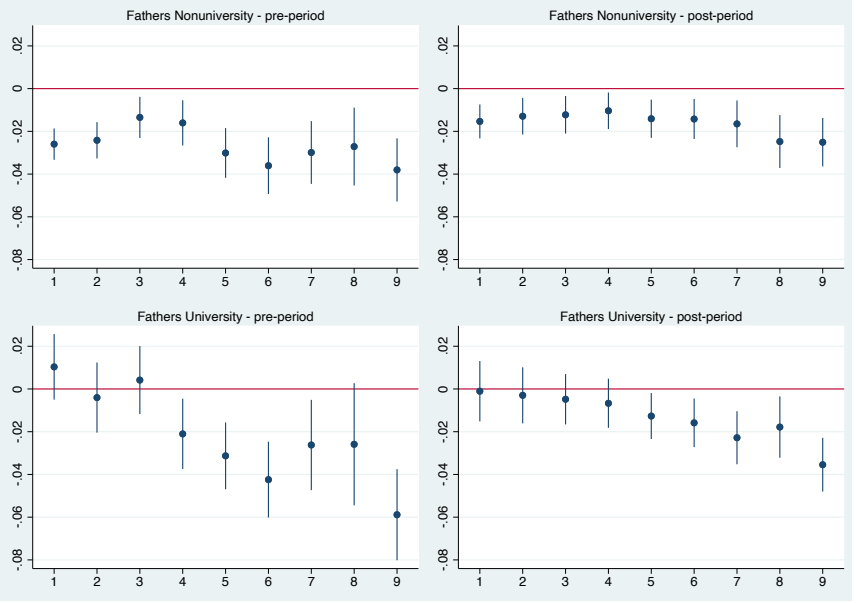
10.2 Wages

Wage Penalties

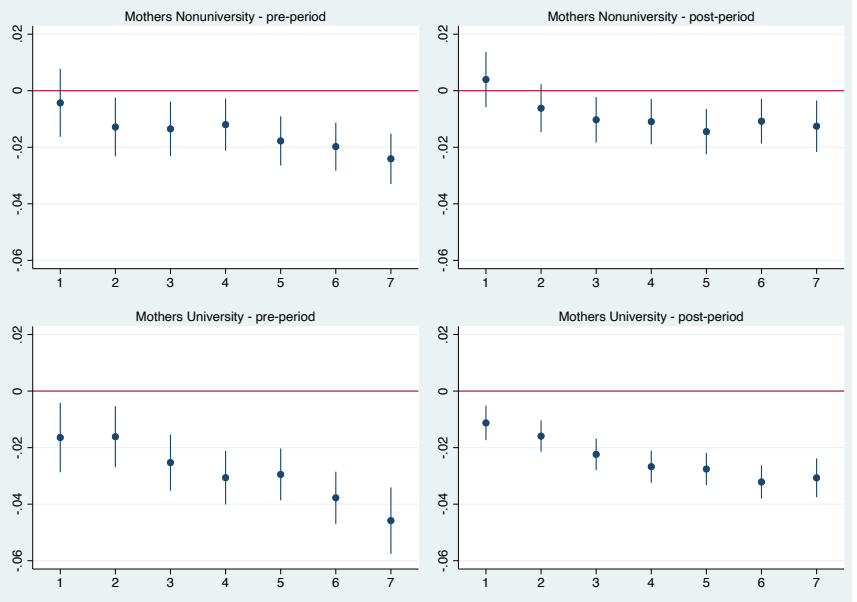
The wage penalties in Table 1 are constructed by averaging coefficients over contiguous intervals in which the point estimates are similar, as described in the text. The graph below shows the estimated coefficients and 95% confidence intervals.

Figure 5: Wage Penalty Estimates

(a) Men



(b) Women



Note: Coefficients from estimation of equation 3. For men, parental leave categories from 1 to 8 represent the number of months taken, and category 9 is for 9 months or more. The omitted category is 0 months. For women, the omitted category is 10 months or less, category 1 is for 11 months, 2 is for 12 months, and so on. For both men and women, the top graphs are for non-university in the pre-period followed by the post period; the bottom graphs are for university in the pre-period followed by post period.

The parameters of the log normal distribution of wages are taken directly from the data. The table below reports the means, variances, and correlation of log wages, expressed in 2014 SEK, from the year prior to the first birth as well as their normalized counterparts for the individuals in the couples sample. We normalize wages by dividing by the mean monthly wage of non-university women in the pre-reform period. (See couples sample Monthly Wage $(t-1)$ in Table A2)

Table A3: Wage Parameters

	Pre-Reform							
	Men				Women			
	nn	nu	un	uu	nn	nu	un	uu
Log wages								
$\mu_w^{g,ij}$	9.840	9.901	10.006	10.038	9.697	9.798	9.770	9.910
$(\sigma_w^{g,ij})^2$	0.513	0.624	0.975	0.933	0.377	0.431	0.470	0.625
Normalized mean wages	1.147	1.226	1.385	1.427	0.987	1.095	1.067	1.234
	Post-Reform							
	Men				Women			
	nn	nu	un	uu	nn	nu	un	uu
Log wages								
$\mu_w^{g,ij}$	10.132	10.195	10.298	10.336	10.005	10.128	10.092	10.212
$(\sigma_w^{g,ij})^2$	1.534	1.638	1.857	1.915	1.342	1.520	1.473	1.667
Normalized mean wages	1.534	1.638	1.857	1.915	1.342	1.520	1.473	1.667
	Pre-Reform				Post-Reform			
	nn	nu	un	uu	nn	nu	un	uu
ρ^{ij}	0.398	0.383	0.388	0.491	0.374	0.384	0.373	0.469

Note: These are the means and variances of real (2014 SEK) log wages for each gender-household type along with the correlation of partners' log wages. All variances on this table are multiplied by 10. We used real log wages from the year prior to the first birth for both the pre-reform period and the post-reform period. We also report normalized mean wages for each group. Specifically, we draw 1,000,000 samples of log wages from its bivariate distribution, exponentiate them to obtain wages, normalize each wage by the mean wage of non-university women in the pre-reform period, and then report the average of these normalized values. Household types are denoted by the education level of the man first and the woman second, where n represents non-university and u represents university education.

Wage Growth Parameters

Table A4 shows the estimates of $\alpha_0^{g,e}$ and other parameters by gender and education.

Table A4: Wage Growth Parameters

	Men							
	Non-Univ				Univ			
Period	$\alpha_0^{g,e}$	$\bar{\phi}^{g,e}$	$\bar{\text{age}}^{g,e}$	Median Age	$\alpha_0^{g,e}$	$\bar{\phi}^{g,e}$	$\bar{\text{age}}^{g,e}$	Median Age
Pre	0.306	0.001	-0.045	30	0.269	-0.005	0.141	32
Post	0.217	-0.004	-0.029	31	0.255	-0.006	-0.022	33
	Women							
	Non-Univ				Univ			
Period	$\alpha_0^{g,e}$	$\bar{\phi}^{g,e}$	$\bar{\text{age}}^{g,e}$	Median Age	$\alpha_0^{g,e}$	$\bar{\phi}^{g,e}$	$\bar{\text{age}}^{g,e}$	Median Age
Pre	0.241	0.001	-0.029	29	0.389	-0.001	-0.087	30
Post	0.182	-0.008	-0.015	28	0.148	-0.005	0.047	31

Note: $\alpha_0^{g,e}$ is the constant term in the wage penalty regression, as shown in equation 3. $\bar{\phi}^{g,e}$ is the average year fixed effect for each gender-education group, and $\bar{\text{age}}^{g,e}$ is the coefficient on the median age dummy for their gender-education group as shown in equation 6.

10.3 The Couple's Maximization Problem

Using the notation in Section 5.2, we are now set to express household income for both the pre- and post-reform periods during the period $\tau = 17$. Note that in the pre-period, $T = 12$ and $D = 1$, whereas in the post-reform period, $T = 13$ and $D = 2$. In both cases, the maximum number of paid months one parent can take is 11, though as a couple they can take more by respecting the daddy month(s).

First, some notation. We denote the wage of the parent with the greater wage by $\bar{w}$ and the wage of the other parent by $\underline{w}$. Similarly, we denote by $\bar{t}$ the parental leave time taken by the parent whose wage is greater and by $\underline{t}$ the parental leave time taken by the parent whose wage is lower. Lastly, we denote by t_d, w_d the parental leave time and wage of a parent who is taking less parental leave than the prescribed daddy months, i.e., $t_d < D$, and by t_{nd}, w_{nd} the parental leave time and wage of the other spouse.[59] Depending on the total length of parental leave time taken by the couple and how this leave is divided between them, we have the following cases for earnings during the observed 17 month period.

One possible case is for both parents to jointly take no more than T months – the period of time in which parental leave is paid – and to respect the daddy months restriction by having each parent take at least D months. In this case, net household income incorporates

[59]It is theoretically possible to have both parents take less than D, in which case the budget constraint is the same as in case 1.

the cost of private child care, C_r, for those months in which both parents are working and the child is below the age of 1, and the public cost C_u when the child is over 1 year old. Total household income is then given by:

- Case 1: $t_m + t_f \leq T,\ \min\{t_g\} \geq D$
  $17\tilde{y} = (17 - t_m)w_m + (17 - t_f)w_f + \kappa(t_m w_m + t_f w_f) - (\max\{12 - t_m - t_f, 0\})C_r - (5 - \max\{t_m + t_f - 12, 0\})C_u$

To understand the algebra in Case 1, note that total household income over the 17 month period is given by the man's monthly wage multiplied by the number of months not spent on parental leave (17 - t_m) (the first term), an equivalent expression for the mother (the second term) plus the parental leave income which is simply the months each parent spent on parental leave multiplied by the replacement rate κ (the third term). From this income, one needs to subtract the cost of the months the child received private childcare, which is positive only if parents are jointly taking less than 12 months of parental leave (hence the max expression). In addition, one needs to subtract the cost of the months in which the child receives public childcare which, given that for this case $t_m + t_f \leq T$, implies that this is either the full remaining 5 months in the pre-reform period since T=12 and either 5 or 4 months depending on whether the couple jointly took 13 months or strictly fewer (when T=13) in the post-reform period for which $D = 2$.[60]

A second case is for the couple to jointly take more than T months, i.e., they take some unpaid parental leave. If the daddy months are respected and no parent takes strictly more than 11 months, this yields the following expression for net household income:

- Case 2: $t_m + t_f > T,\ \min\{t_g\} \geq D$ and $t^M \equiv \max\{t_g\} \leq 11$
  $17\tilde{y} = (17 - t_m)w_m + (17 - t_f)w_f + \kappa(\bar{w}\bar{t} + \underline{w}(T - \bar{t})) - (17 - t_m - t_f)C_u$

Note that in this case the parent with the greater monthly wage will take all their parental leave months as paid whereas the other parent will have some months unpaid.

A third case is that the couple jointly takes more than the paid parental leave but, in contrast to case 2, one parent takes more than 11 months. Assuming that the daddy months

[60]The algebra for the remaining cases can be derived similarly so we do not provide detailed explanations for them.

are taken, we can distinguish between two sub-cases according to whether the parent with the higher wage takes more or less than 11 months:

- Case 3: $t_m + t_f > T,\ \min\{t_g\} \geq D$ *and* $t^M \equiv \max\{t_g\} > 11$
  - 3a). $\bar{t} > 11$
    $17\tilde{y} = (17 - t_m)w_m + (17 - t_f)w_f + \kappa(11\bar{w} + D\underline{w}) - (17 - t_m - t_f)C_u$
  - 3b). $\bar{t} \leq 11$
    $17\ \tilde{y} = (17 - t_m)w_m + (17 - t_f)w_f + \kappa(\bar{t}\bar{w} + \underline{w}(T - \bar{t})) - (17 - t_m - t_f)C_u$

Lastly, suppose that the daddy months are not fully taken. Here we can distinguish two cases according to whether the other parent takes more or less than 11 months (the amount that is paid).

- Case 4: $\min\{t_g\} < D$
  - 4a). $t_{nd} \leq 11$
    $17\ \tilde{y} = (17 - t_m)w_m + (17 - t_f)w_f + \kappa(t_d w_d + t_{nd} w_{nd}) - (12 - t_m - t_f)C_r - 5C_u$
  - 4b). $t_{nd} > 11$
    $17\tilde{y} = (17 - t_m)w_m + (17 - t_f)w_f + \kappa(t_d w_d + 11 w_{nd}) - (17 - t_m - t_f)C_u$

10.4 Estimation Procedure

We start the estimation process with an initial guess of the parameter vector θ_0. For this guess of parameters, we solve for the steady states for both periods sequentially: first solving for the pre-reform period and then for the post-reform period. To begin, we use the pre-reform empirical means of parental leave for each gender and household type, denoted as $\bar{t}_0^{g,ij}$, as our starting point. Note that this choice then determines the means of the γ distributions ($\mu_{\gamma,0}^{g,ij}$) as $a^{g,e}$ and $b^{g,e}$ are specified in θ_0. We then simulate the parental leave decisions of 20,000 households (5,000 for each household type) under the prevailing budget constraints, wages, and wage penalties of the pre-reform environment. This simulation produces a distribution of parental leave taken for each gender and household type, from which we calculate updated means of parental leave, $\bar{t}_1^{g,ij}$, for each gender-household type.

Next, we check for convergence to a fixed point by comparing the differences between these new means ($\bar{t}_1^{g,ij}$) and initial means ($\bar{t}_0^{g,ij}$). Specifically, we evaluate if the maximum absolute difference in means across all groups between $\bar{t}_1^{g,ij}$ and $\bar{t}_0^{g,ij}$ is less than our convergence threshold of 0.01. If the convergence criterion is not met, $\bar{t}_1^{g,ij}$ are used to simulate another 20,000 households with updated $\mu_{\gamma,1}^{g,ij}$ parameters, yielding a new set of parental leave means, $\bar{t}_2^{g,ij}$. The process of calculating the differences of parental leave means between successive iterations continues until convergence between successive means ($\bar{t}_{s-1}^{g,ij}$ and $\bar{t}_s^{g,ij}$) is achieved. Upon reaching a fixed point, we perturb the resulting means (for all gender and household types) by a small amount, ϵ, to ensure the model returns to the same mean values, thus verifying local stability.[61]

After solving for the steady state of the pre-reform period, we employ the same process to solve for the post-reform period steady state, with the same guess of the parameter vector θ_0 but with the post-reform opportunity set (i.e, a second daddy month) and the post-reform budget constraints, wages, and wage penalties. For the post-reform simulations, however, we use the previously obtained steady-state parental leave means as the starting point to define $\mu_{\gamma,0}^{g,ij}$. The convergence and local stability criteria remain the same as in the pre-reform period.

Once steady states for both periods are found, we generate moments for both pre- and post-reform periods using the steady-state distributions, and compare them to the empirical moments. We then update our guess for the parameter vector to θ_1 using the Nelder-Mead algorithm and repeat the process of finding steady states and computing moments for both periods with a new guess of parameters.[62] This iterative process continues until we find a parameter vector that minimizes our objective function.

10.5 Derivative Analysis

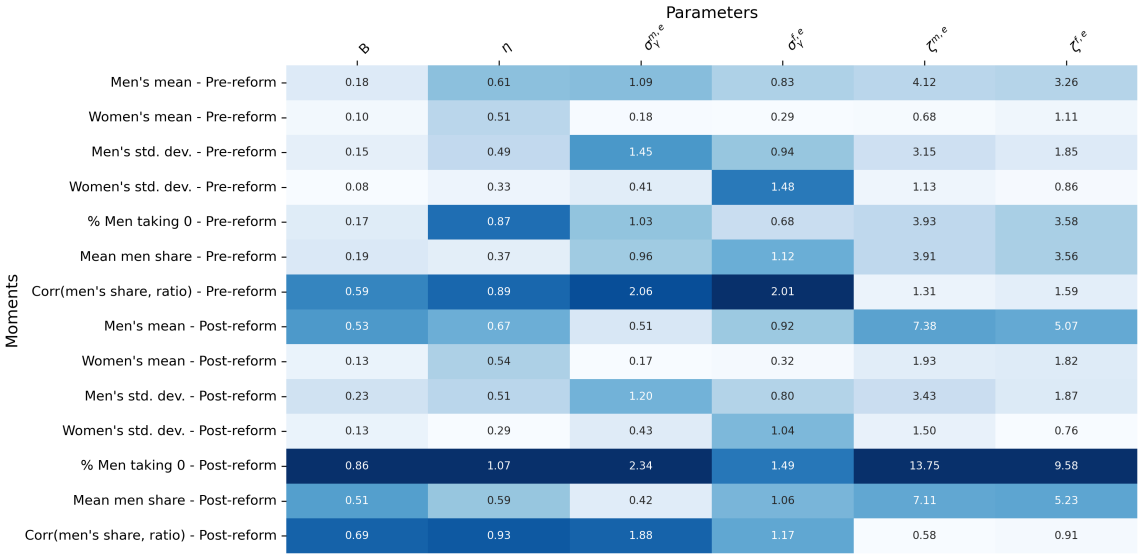
In Figure 6, each parameter (column) is color-coded to indicate the relative magnitude of a 1% perturbation's effect on the moments, with darker shading corresponding to larger effects and lighter shading reflecting smaller ones. Although estimation moments are defined

[61] We also checked global stability, which was also satisfied but we do not have a general proof.

[62] The Nelder-Mead algorithm was chosen for its robustness in handling non-differentiable models typical of discrete choice frameworks.

by household types, they are averaged into a single measure for simplicity of presentation. Similarly, parameters have been aggregated by gender where appropriate to underscore overall patterns rather than subgroup-level variation.

Figure 6: Derivative Analysis



10.6 Counterfactual Correlation

Here are the results of targeting a counterfactual uniform correlation between the household male to female wage ratio and the man's share of leave for all household types of -0.60. As can be seen from the table, the model does less well in matching several moments related to the length of women's parental leave.[63]

[63]The sum of squared errors is considerably higher - 1.7 as contrasted with 0.7.

Table A5: Alternative Correlation: From Pre to Post Reform

Moments	Pre-Reform Simulated Moments	Reform Only	Income	Social Norms	Post-Reform Simulated Moments
			nn		
Mean months of men	1.82	1.88	2.24	1.92	2.55
SD months of men	2.19	2.21	3.10	2.22	3.20
Mean months of women	9.91	10.03	9.50	10.00	9.26
SD months of women	2.00	1.96	2.93	1.98	3.07
% Men taking 0 month	27.47	27.47	26.68	26.57	19.39
Mean share taken by men	15.16	15.30	18.63	15.60	21.26
Corr between wage ratio & men's share	-0.40	-0.41	-0.53	-0.41	-0.58
			nu		
Mean months of men	2.21	2.26	2.93	2.30	3.43
SD months of men	2.41	2.42	3.42	2.43	3.46
Mean months of women	9.55	9.61	8.85	9.58	8.45
SD months of women	2.26	2.25	3.26	2.27	3.37
% Men taking 0 month	23.29	23.36	22.23	22.51	11.48
Mean share taken by men	18.50	18.70	24.40	19.01	28.62
Corr between wage ratio & men's share	-0.39	-0.39	-0.57	-0.40	-0.64
			un		
Mean months of men	2.12	2.16	2.53	2.20	2.73
SD months of men	2.45	2.51	3.26	2.52	3.32
Mean months of women	9.64	9.74	9.28	9.72	9.12
SD months of women	2.28	2.22	3.14	2.23	3.22
% Men taking 0 month	24.02	24.05	18.77	23.29	14.45
Mean share taken by men	17.68	17.61	21.08	17.88	22.77
Corr between wage ratio & men's share	-0.43	-0.43	-0.55	-0.43	-0.57
			uu		
Mean months of men	2.74	2.78	3.24	2.81	3.46
SD months of men	2.58	2.61	3.41	2.62	3.42
Mean months of women	9.08	9.16	8.63	9.14	8.45
SD months of women	2.45	2.43	3.31	2.44	3.35
% Men taking 0 month	17.01	17.09	12.50	16.53	8.62
Mean share taken by men	22.88	22.88	27.05	23.11	28.87
Corr between wage ratio & men's share	-0.47	-0.47	-0.65	-0.47	-0.67

Note: This exercise uses re-estimated parameters from setting the correlation between the male to female household wage ratio and the within-household share of leave taken by men to -0.6 across all household types. Column 2 introduces the reform but no other changes. Column 3 allows, in addition to the reform, all income related variables (wage parameters, wage penalties, and childcare costs) to take their post-period values as given in Tables 1, 2, and A4. Social norms (the $\mu_\gamma^{g,ij}$) are kept constant at the pre-reform values. Column 4 returns to the wages and wage penalties of the pre-reform period but now allows the values of the $\mu_\gamma^{g,ij}$ to evolve endogenously. Column 5 allows all changes. Household types are denoted by the education level of the man first and the woman second, where n represents non-university and u represents university education.

A Finer Decomposition of Income-Related Changes

Table A6: Income-Related Changes: From Pre to Post Reform

Moments	Pre-Reform Simulated Moments	Reform Only	Reform + Wage Dist	Reform + Childcare Cost	Reform + Wage Penalty	Reform + All Income	Social Norms	Post-Reform Simulated Moments
				nn				
Mean months of men	2.05	2.25	2.26	2.23	2.34	2.33	2.64	2.99
SD months of men	2.03	2.01	2.03	1.99	2.12	2.11	2.15	2.31
Mean months of women	13.48	13.42	13.38	13.31	13.49	13.40	13.04	12.77
SD months of women	2.80	2.63	2.65	2.62	2.69	2.71	2.71	2.83
% Men taking 0 month	27.28	27.26	27.29	27.29	27.05	27.06	21.39	17.86
Mean share taken by men	13.35	14.41	14.46	14.41	14.80	14.81	16.91	19.08
Corr(wage ratio, men's share)	-0.07	-0.07	-0.07	-0.07	-0.06	-0.06	-0.07	-0.06
				nu				
Mean months of men	2.54	2.74	2.76	2.72	2.85	2.84	3.17	3.75
SD months of men	2.25	2.22	2.24	2.19	2.31	2.30	2.31	2.43
Mean months of women	12.53	12.59	12.54	12.48	12.69	12.56	12.28	11.85
SD months of women	2.82	2.69	2.70	2.68	2.84	2.82	2.73	2.84
% Men taking 0 month	22.32	22.34	22.29	22.41	22.07	22.05	16.07	11.50
Mean share taken by men	16.85	17.84	17.97	17.84	18.34	18.42	20.50	24.07
Corr(wage ratio, men's share)	-0.09	-0.10	-0.10	-0.09	-0.09	-0.09	-0.10	-0.10
				un				
Mean months of men	2.30	2.44	2.46	2.43	2.57	2.58	2.70	3.51
SD months of men	2.24	2.21	2.24	2.20	2.29	2.30	2.32	2.62
Mean months of women	13.53	13.50	13.45	13.42	13.43	13.33	13.25	12.37
SD months of women	2.78	2.64	2.66	2.64	2.73	2.75	2.71	2.96
% Men taking 0 month	25.61	25.62	25.64	25.63	25.86	25.88	21.84	15.46
Mean share taken by men	14.58	15.28	15.42	15.29	16.09	16.21	16.90	22.14
Corr(wage ratio, men's share)	-0.08	-0.08	-0.08	-0.08	-0.08	-0.08	-0.08	-0.08
				uu				
Mean months of men	3.14	3.27	3.29	3.26	3.45	3.44	3.62	4.61
SD months of men	2.65	2.60	2.62	2.59	2.62	2.62	2.70	2.83
Mean months of women	12.26	12.33	12.35	12.25	12.25	12.22	12.07	11.26
SD months of women	2.93	2.81	2.83	2.80	2.90	2.91	2.84	2.96
% Men taking 0 month	17.49	17.56	17.59	17.57	17.64	17.69	14.08	8.77
Mean share taken by men	20.18	20.74	20.76	20.77	21.84	21.81	22.84	28.90
Corr(wage ratio, men's share)	-0.10	-0.10	-0.10	-0.10	-0.10	-0.09	-0.11	-0.10

Note: Column 2 introduces the reform but no other changes. Column 3, in addition to the reform, adopts post-period wage parameters from Table A4. Column 4 instead adopts post-period childcare costs from Table 2. Column 5 adopts post-period wage penalties from Table 1. Column 6 allows all income-related variables (wage parameters, wage penalties, and childcare costs) to take their post-period values. In Columns 3–6, social norms (the $\mu_\gamma^{g,ij}$) remain fixed at their pre-reform levels. Column 7 returns income-related variables to their pre-reform values but lets the social-norm parameters $\mu_\gamma^{g,ij}$ evolve endogenously. Column 8 incorporates all changes simultaneously. Household types are denoted by the education level of the man first and the woman second, where n indicates non-university and u indicates university education.

10.7 Regression Discontinuity

As described in the text, We exploit the discontinuity introduced by the 2002 "Daddy Month" reform to estimate its impact on parental leave (PL) uptake using a sharp regression discontinuity (RD) design. The outcome is measured as the number of months of leave taken within the first 16 months following childbirth. Our analysis is based on the couples sample

and includes births in 2001 and 2002. The cutoff is defined at January 2002, and a monthly running variable is used. The sample consists of 6,059 couples, with monthly observations ranging from approximately 200 to 300 couples.

We estimate the discontinuity at the cutoff (January 2002) separately for fathers and mothers. As noted in the text, the RD equation estimated is given by:

$$PL_i = \alpha + \delta D_i + f(X_i) + u_i \tag{9}$$

where PL_i is the number of parental leave months taken within the 16 months following childbirth. The indicator variable D_i equals 1 if the child was born in 2002 and 0 if the child was born in 2001. The running variable X_i ranges from -12 to 11, taking the value 0 at the cutoff (January 2002). It represents months relative to the reform date, with negative values for births in 2001 and positive values for births in 2002.

We apply a triangular kernel and manually select bandwidths of 12 months on each side of the cutoff. Both mean (order $p = 0$) and linear (order $p = 1$) polynomial specifications of $f(X)$ are estimated to assess the robustness of the results to the choice of functional form. Standard errors are computed using the nearest-neighbor variance estimation method.

The RD plots are presented in Appendix Figure 7. The first panel shows results for fathers, including a polynomial fit of order 0 (mean) and 1 (linear), respectively. As discussed in the text, there appears to be a discontinuity in fathers' parental leave uptake of slightly less than half a month. In contrast, for mothers, there is no clear evidence of a discontinuity associated with the 2002 reform.

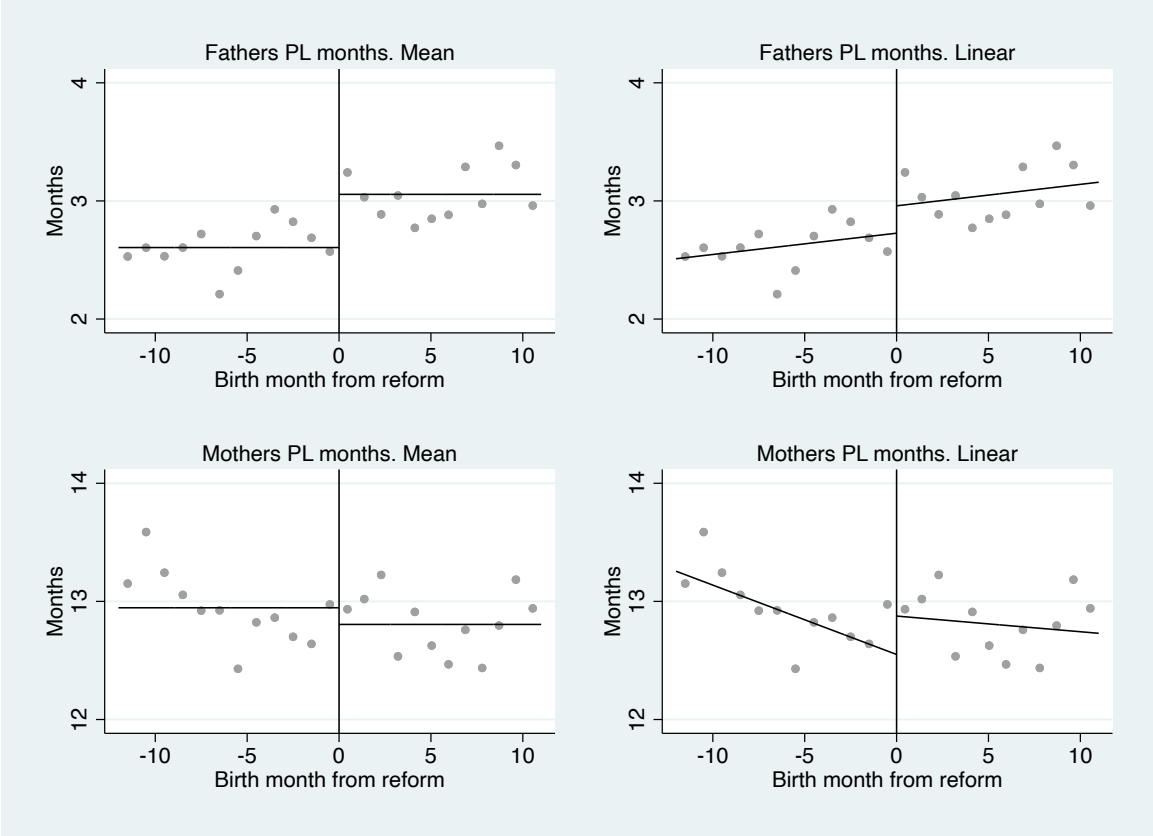
The RD estimates are reported in Table A7. For fathers, we observe a positive and statistically significant discontinuity at the cutoff under the mean specification ($p = 0$), with an estimated effect of 0.388 months (SE = 0.076). The corresponding linear specification ($p = 1$) yields a somewhat smaller effect of 0.268 months (SE = 0.154).

For mothers, the estimated discontinuities are smaller and statistically insignificant in both specifications. The mean polynomial produces an estimated effect of -0.014 months (SE = 0.099), while the linear specification yields an estimate of 0.316 months (SE = 0.203). The larger standard errors for mothers suggest greater uncertainty around these estimates.

Overall, while we do not claim to provide a causal analysis, the results indicate a clearer

and more precisely estimated RD effect for fathers, consistent with the direct targeting of the reform towards increasing paternal leave uptake. No significant discontinuity is detected for mothers.

Figure 7: RD Plots of Monthly Parental Leave



Note: The figure displays RD plots of monthly parental leave uptake for fathers and mothers separately. The running variable is measured in months relative to the January 2002 cutoff. The figures are based on the couples sample.

The table below reports the 2001 pre-reform log-wage distribution parameters for men and women (μ_f^{hh}, σ_f^{hh}) by household type along withthe number of observations in the RD sample both in 2001 (pre-reform) and 2002 (post-reform). Panel C shows the corresponding household-type shares (weights) used to aggregate the model's household-type–specific means into single gender-specific means.

Table A7: Regression Discontinuity Estimates of the 2002 "Daddy Month" Reform on Parental Leave Usage within 16 Months Following Childbirth

	PL Father	PL Father	PL Mother	PL Mother
RD_Estimate	0.388***	0.268*	-0.014	0.316
	(0.076)	(0.154)	(0.099)	(0.203)
Observations	6,059	6,059	6,059	6,059
Polynomial	Mean	Linear	Mean	Linear
Bandwidth	12 months	12 months	12 months	12 months

Note: The table reports RD estimates of parental leave uptake within 16 months of childbirth for fathers and mothers separately. A triangular kernel and a bandwidth of 12 months on each side of the cutoff are used. Both mean (order $p = 0$) and linear (order $p = 1$) polynomial specifications are presented. Standard errors are shown in parentheses. The estimations are based on the couple sample.

Table A8: RD Parameters for Model Simulation

Panel A: Pre-Reform 2001
Log Wage Parameters

	nn	**nu**	**un**	**uu**
Men				
$\mu_w^{m,ij}$	9.91	9.96	10.09	10.09
$\sigma_w^{m,ij}$	0.24	0.26	0.33	0.31
Women				
$\mu_w^{f,ij}$	9.77	9.85	9.84	9.96
$\sigma_w^{f,ij}$	0.21	0.20	0.24	0.25

Panel B: RD Sample Observations

	nn	**nu**	**un**	**uu**
Pre-Reform, 2001	1390	571	223	678
Post-Reform, 2002	1384	668	302	843

Panel C: Weights

	nn	**nu**	**un**	**uu**
Pre-Reform, 2001	0.49	0.20	0.08	0.24
Post-Reform, 2002	0.43	0.21	0.09	0.26

Note: Panel A reports the 2001 pre-reform log-wage distribution parameters for men ($\mu_w^{m,ij}$, $\sigma_w^{m,ij}$) and women ($\mu_w^{f,ij}$, $\sigma_w^{f,ij}$) by household type. Panel B gives the number of observations in the RD sample for each household type in 2001 (pre-reform) and 2002 (post-reform). Panel C shows the corresponding household-type shares (weights) used to aggregate the model's household-type–specific means into single gender-specific means. Household types are denoted by the education level of the man first and the woman second, where n represents non-university and u represents university education.

References

Albrecht, James W., Per-Anders Edin, Marianne Sundström and Susan B. Vroman. 1999. "Career Interruptions and Subsequent Earnings: A Reexamination Using Swedish Data." *The Journal of Human Resources* 34(2):294–311.

Angelov, Nikolay, Per Johansson and Erica Lindahl. 2016. "Parenthood and the Gender Gap in Pay." *Journal of Labor Economics* 34(3):545–579.

Avdic, Daniel and Arizo Karimi. 2018. "Modern Family? Paternity Leave and Marital Stability." *American Economic Journal: Applied Economics* 10(4):283–307.

Canaan, Serena, Anne Lassen, Philip Rosenbaum and Herdis Steingrimsdottir. 2022. "Maternity leave and paternity leave: Evidence on the economic impact of legislative changes in high income countries." *IZA Discussion Paper* .

Carlsson, Magnus and Abdulaziz Abrar Reshid. 2022. "Co-worker peer effects on parental leave take-up." *The Scandinavian Journal of Economics* 124(4):930–957.
URL: *https://onlinelibrary.wiley.com/doi/abs/10.1111/sjoe.12485*

Carneiro, Pedro, Katrine V. Løken and Kjell G. Salvanes. 2015. "A Flying Start? Maternity Leave Benefits and Long-Run Outcomes of Children." *Journal of Political Economy* 123(2):365–412.

Chiplunkar, Gaurav and Pinelopi Koujianou Goldberg. 2024. "Aggregate Implications of Barriers to Female Entrepreneurship." *Econometrica* 92(6):1801–1835.

Cools, Sara, Jon H. Fiva and Lars J. Kirkebøen. 2015. "Causal Effects of Paternity Leave on Children and Parents." *The Scandinavian Journal of Economics* 117(3):801–828.

Dahl, Gordon B, Katrine V Løken and Magne Mogstad. 2014. "Peer effects in program participation." *American Economic Review* 104(7):2049–2074.

Dahl, Gordon B., Katrine V. Løken, Magne Mogstad and Kari Vea Salvanes. 2016. "What Is the Case for Paid Maternity Leave?" *The Review of Economics and Statistics* 98(4):655–670.

Dottori, Davide, Francesca Modena and Giulia Martina Tanzi. 2024. "Peer effects in parental leave: Evidence from Italy." *Labour Economics* .

Dustmann, Christian and Uta Schönberg. 2012. "Expansions in maternity leave coverage and children's long-term outcomes." *American Economic Journal: Applied Economics* 4(3):190–224.

Duvander, Ann-Zofie and Mats Johansson. 2012. "What are the effects of reforms promoting fathers' parental leave use?" *Journal of European Social Policy* 22(3):319–330.

Duvander, Ann-Zofie and Mats Johansson. 2019. "Does Fathers' Care Spill Over? Evaluating Reforms in the Swedish Parental Leave Program." *Feminist Economics* 25(2):67–89.

Duvander, Ann-Zofie and Sofie Cedstrand. 2022. 263Gender Equal Parental Leave Use in Sweden: The Success of the Reserved Months. In *Successful Public Policy in the Nordic Countries: Cases, Lessons, Challenges*. Oxford University Press.

Ekberg, John, Rickard Eriksson and Guido Friebel. 2013. "Parental leave—A policy evaluation of the Swedish "Daddy-Month" reform." *Journal of Public Economics* 97:131–143.

Erosa, Andrés, Luisa Fuster and Diego Restuccia. 2010. "A general equilibrium analysis of parental leave policies." *Review of Economic Dynamics* 13(4):742–758.

Farré, L, C. Felfe, L. Gonzalez and P. Schneider. 2023. Changing Gender Norms Across Generations: Evidence from a Paternity Leave Reform. Working Paper 16341 IZA.

Farré, Lídia and Libertad González. 2019. "Does paternity leave reduce fertility?" *Journal of Public Economics* 172:52–66.

Fernández, Raquel. 2013. "Cultural Change as Learning: The Evolution of Female Labor Force Participation over a Century." *American Economic Review* 103(1):472–500.

Fernández, Raquel. 2025. "Understanding Cultural Change." work in progress for a Chapter in the Handbook of Culture and Economic Behavior, edited by Enke, B and P. Giuliano and N. Nunn and L. Wantchekon, Elsevier.

Försäkringskassann. 2024. "The history of parental insurance." https://www.forsakringskassan.se/privatperson/foralder/foraldraforsakringens-historia. Accessed: July 25, 2024.

Jørgensen, Thomas H. and Jakob E. Søgaard. 2024. "The division of parental leave: Empirical evidence and policy design." *Journal of Public Economics* 238:105202.

Kleven, Henrik, Camille Landais and Jakob Egholt Søgaard. 2019. "Children and Gender Inequality: Evidence from Denmark." *American Economic Journal: Applied Economics* 11(4):181–209.

Kotsadam, Andreas and Henning Finseraas. 2011. "The state intervenes in the battle of the sexes: Causal effects of paternity leave." *Social Science Research* 40(6):1611–1622.

Kotsadam, Andreas and Henning Finseraas. 2013. "Causal Effects of Parental Leave on Adolescents' Household Work." *Social Forces* 92(1):329–351.

Lalive, Rafael and Josef Zweimüller. 2009. "How Does Parental Leave Affect Fertility and Return to Work? Evidence from Two Natural Experiments*." *The Quarterly Journal of Economics* 124(3):1363–1402.

Linderoth, Anna. 2024. "Reference Points for Men's Parental Leave-Taking Behavior: Evidence from Swedish Reforms." Working paper.

Moberg, Ylva and Maaike van der Vleuten. 2022. "Why do gendered divisions of labour persist? Parental leave take-up among adoptive and biological parents." *European Sociological Review* 39(2):210–228.

Olivetti, Claudia and Barbara Petrongolo. 2017. "The Economic Consequences of Family Policies: Lessons from a Century of Legislation in High-Income Countries." *Journal of Economic Perspectives* 31(1):205–30.

Rossin-Slater, Maya. 2018. Maternity and Family Leave Policy. In *The Oxford Handbook of Women and the Economy*. Oxford University Press.

Ruhm, Christopher J. 1998. "The Economic Consequences of Parental Leave Mandates: Lessons from Europe*." *The Quarterly Journal of Economics* 113(1):285–317.

Saez, Emmanuel. 2010. "Do Taxpayers Bunch at Kink Points?" *American Economic Journal: Economic Policy* 2(3):180–212.

Schönberg, Uta and Johannes Ludsteck. 2014. "Expansions in Maternity Leave Coverage and Mothers' Labor Market Outcomes after Childbirth." *Journal of Labor Economics* 32(3):469–505.

Skolverket. 2000. "Beskrivande data om barnsomorg och skola 1999.". Rapport 192.

Skolverket. 2005. "Beskrivande data 2005. Förskolverksamhet, skolbarnsomsorg, skola och vuxenutbildning.". Rapport 265.

Skolverket. 2010. ""Kostnader för förskoleverksamhet, skolbarnsomsorg, skola och vuxenutbildning 2009".". PM Dnr 71-2010:4.

Tervola, Jussi, Ann-Zofie Duvander and Eleonora Mussino. 2017. "Promoting Parental Leave for Immigrant Fathers—What Role Does Policy Play?" *Social Politics: International Studies in Gender, State Society* 24(3):269–297.

Tô, Linh T. 2018. "The Signaling Role of Parental Leave." *Mimeo* .

Welteke, Clara and Katharina Wrohlich. 2019. "Peer effects in parental leave decisions." *Labour Economics* 57:146–163.

World Economic Forum, Report. 2024. "Global Gender Gap Report 2024.". https://www.weforum.org/publications/global-gender-gap-report-2024/.

Yamaguchi, Shintaro. 2019. "Effects of parental leave policies on female career and fertility choices." *Quantitative Economics* 10(3):1195–1232.