ROCKWOOL Foundation Berlin
Institute for the Economy and the Future of Work (RFBerlin)

DISCUSSION PAPER SERIES 038/26

Parental Leave: Economic Incentives and Cultural Change

Jim Albrecht, Per-Anders Edin, Raquel Fernandez, Jiwon Lee, Peter Skogman Thoursie, Susan Vroman

www.rfberlin.com

JANUARY 2026

Parental Leave: Economic Incentives and Cultural Change

Authors

Jim Albrecht, Per-Anders Edin, Raquel Fernandez, Jiwon Lee, Peter Skogman Thoursie, Susan Vroman

Reference

JEL Codes: D10, J16, Z10, Z18

Keywords: Parental Leave, Gender Equality, Childcare, Culture

Recommended Citation: Jim Albrecht, Per-Anders Edin, Raquel Fernandez, Jiwon Lee, Peter Skogman Thoursie, Susan Vroman (2026): Parental Leave: Economic Incentives and Cultural Change. RFBerlin Discussion Paper No. 038/26

Access

Disclaimer

Imprint

RFBerlin
ROCKWOOL Foundation Berlin –
Institute for the Economy
and the Future of Work

Gormannstrasse 22, 10119 Berlin
Tel: +49 (0) 151 143 444 67
E-mail: info@rfberlin.com
Web: www.rfberlin.com

Parental Leave:
Economic Incentives and Cultural Change*

James Albrecht
Georgetown Univ., IZA, RFBerlin

Per-Anders Edin
Uppsala Univ., IFAU

Raquel Fernández
NYU, NBER, CEPR, IZA, RFBerlin

Jiwon Lee
NYU

Peter Skogman Thoursie
Stockholm Univ.

Susan Vroman
Georgetown Univ., IZA, RFBerlin

January 2026

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 substantially larger 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

*We thank the following people for helpful comments: Malin Tallås Ahlzén, Milena Almagro, Neha Bairoliya, Diego Daruich, Thomas Jørgensen, Andreas Kotsadam, and seminar audiences at the Di Tella Conference, the Gender in the Global Economy Conference, the Narratives and Culture in Historical Perspective Conference, the NBER SI Gender in the Economy Conference, Carlos III University, IFAU Uppsala, University of Houston, St. Louis Fed., Duke University, UCLA, USC, Uppsala University, IIES Stockholm, SITE Gender 2025, University of Delaware, Human Capital in Macroeconomics Conference, LACEA 2025, CAF-WELAC Gender Conference, and ASREC at the ASSA 2026. Albrecht and Vroman thank IFAU for financial support and Fernández thanks the CV Starr Center. Peter Skogman Thoursie acknowledges financial support from the Jan Wallanders och Tom Hedelius Foundation and the Tore Browaldh Foundation.

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 was an early adopter 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 (along with an additional earnings-related month) and a third one in 2016.[4]

The objective of our paper is to use the second-daddy-month reform to study the role of cultural change versus more traditional economic factors in determining the change in parental leave uptake. Doing so also allows us to study how future policies may affect the gender distribution of parental leave. This reform – as opposed to the one-daddy month reform that preceded it – is particularly interesting because it also increased the total number of earnings-related months of parental leave, allowing mothers to still take 11 months as long as fathers take two months.

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

[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 the 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 Försäkringskassan (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.

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, 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.[5]

To understand and quantify how these forces interact, we develop a simple 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 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 household type.

We estimate our model using Swedish administrative data primarily from two time periods: the pre-reform period of 1998-2001 and the post-reform period of 2008-2011. We take (pre-birth) wages in each period as given and estimate the future wage penalty

[5]Throughout we will refer to social norms and culture interchangeably. Furthermore, we do not differentiate between peer effects and culture in the sense that if one's peers take up a particular activity, for example, then one is more likely to also take up that activity. This is how cultural change often happens. See Fernández (2025) for an in depth discussion.

associated with different lengths of parental leave for men and women separately by education, 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 simulated 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 policy experiments, 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 eliminating the gender wage gap increases men's share of parental leave but still leaves it below 30% for all but households in which both parents are university educated and, even for these, men's share is 32.7%. Perhaps surprisingly, decreasing the cost of childcare has very little effect on men's share. These two exercises show that economic reforms may not lead to large changes in parental leave behavior even if culture responds endogenously. The last two policies which radically change parental endowments, on the other hand, significantly increase men's share of parental leave and are much closer to achieving gender parity.

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 relatively small effects, but in our model culture is determined by what other people do rather than by government policy directly and hence evolves endogenously.[6]

There is also a large and growing literature on the effects of parental leave. These papers have focused on a diverse set of issues ranging from the consequences of parental leave for maternal welfare, marital stability, and children's health and human capital (e.g., Avdic and Karimi (2018), 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),Gonzalez and Zoabi (2023) 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).[7] 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 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 significantly increased the subsequent uptake 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

[6] 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.

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

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 age six. The authors find that the reform increased the take-up of parental leave for mothers especially if the share of post-reform peers taking leave was larger. Our paper thinks of these peer effects as contributors to cultural change. In our structural model, peers are more generally defined as individuals of the same gender and household type and we too find strong peer effects.

Our paper also contributes to the literature on cultural change, especially regarding gender roles.[8] 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 us 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

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

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, preferences themselves 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 leave system and the time periods of the analysis. Section 3 describes the datasets, the sample construction, and the distribution of parental leave take-up. Section 4 presents the model. Section 5 presents the model's estimation and discusses identification. Section 6 disentangles the contributions of the reform, of income-related changes, and of changes in culture or social norms to the take-up of parental leave by gender and household type. Section 7 conducts a regression discontinuity analysis of the reform and contrasts these results with those obtained from simulating the model. It also shows that our results regarding the primacy of culture depend on features such as the correlation of the household wage ratio to the share of parental leave taken by the father. Section 8 studies several potential alternative policies and the role of the gender wage gap. Section 9 concludes.

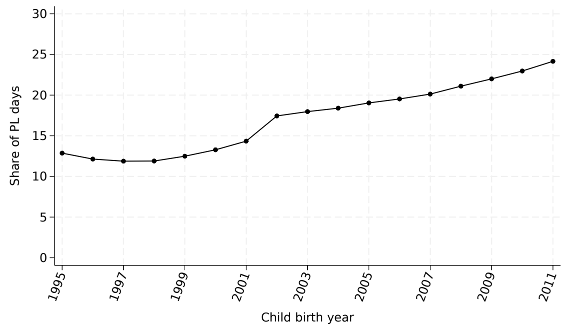
2 Some Preliminaries

2.1 The Parental Leave System

Sweden has been a leader in parental leave. In 1974, it became the first country to extend paid parental leave equally to fathers.[9] 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

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

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.

themselves of this option.[10] 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).[11] Furthermore, a parent's job was protected for up to 18 months.[12]

An important change in Sweden's parental leave system was the introduction of months reserved for each parent – an innovation that Norway spearheaded in 1993 and that later spread to other countries, particularly in the European Union. The first "daddy month" was introduced in Sweden in 1995. In principle, the reform allocated one of the earnings-related months of benefits to each of the parents. The remaining 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.[13] 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.

[10]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)).

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

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

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

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% but check robustness to lower rates. 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.[14]

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. As noted above, 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 1995 reform for several reasons: i. an increase in the intensive margin of reserved months is interesting from a cultural perspective; ii. several papers have already focused on the 1995 reform and, most importantly, iii. the data provided by the Social Insurance Agency is incomplete for 1994, the first year in which parental leave data from the agency is available.

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

[14] 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.

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

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 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.

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.[16] We use the start and end dates for parental leave spells to construct a measure of the months an individual spent 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.[17] 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

[16]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.

[17]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.

option to start taking paid leave 2 months prior to the 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.[18] Restricting our window to 16 months after birth avoids confounding parental leave associated with the first birth with parental leave associated with subsequent births.[19] 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. 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 parental leave $\in (t-1, t]$.

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.[20] 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 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 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.

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

[19] Only 4% of couples have an additional child within 16 months.

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

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. Furthermore, the cultural attitudes of immigrants are very likely to differ from those of native Swedes.[21] Second, we limit our sample to heterosexual couples between the ages of 20 and 60 at the time of childbirth since we are interested in the gender division of 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 required in order to think about the consequences of different lengths of leave. 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.

Eliminating parents with missing wage observations in those years substantially reduces our sample. When we drop mothers for whom we have no wage observation in $t-1$, we are left with 54,337 in the pre-period and 69,823 in the post-period. Imposing the requirement that wages are observed in both $t-1$ and $t+3$ leaves us with 29,178 in the pre-period and 39,176 in the post-period.[22] For fathers, the drop due to a lack of a wage observation in $t-1$ is larger. We have 40,667 in the pre-period and 52,093 in the post-period.

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

[22] 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.

Imposing the requirement that wages are observed in both $t-1$ and $t+3$ for fathers yields 28,285 observations in the pre-period and 36,429 in the post-period. Note that we include individuals even if we do not have wage data for their partners. We refer to this as the "Wage sample."

As one can see, there is a large drop in observations when we move to the Wage sample. This is consistent with the institutional design of the Wage Structure statistics. The public sector accounts for approximately 35 percent of total employment in Sweden and is fully covered by the Wage Structure survey during the reference month. In contrast, the survey is designed to cover roughly 50 percent of private-sector employees during the same month. Conditional on being employed in the reference month, the implied coverage rate is therefore approximately 67.5 percent $(0.35 \times 1 + 0.65 \times 0.5)$.[23]

Average monthly wages are reported in Table A1 expressed in 2014 SEK. There was substantial real wage growth from the pre-period to the post-period. Wages for both university (those with at least 3 years of university, as is the norm in Sweden) and non-university men rose about 35% as did those of non-university women, while the wages of university women rose about 38% .

The Couples Sample

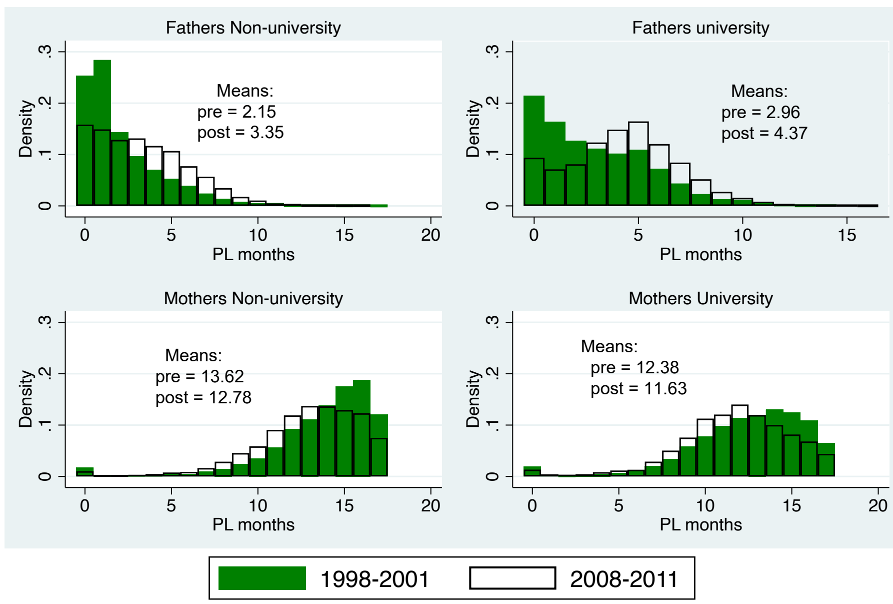
We take the couple as our unit of observation for parental leave, as a leave decision made by one parent necessarily affects the other. This requires us to further restrict our sample to observations for which we have wages for *both* parents, reducing the sample to 9,926 couples in the pre-period and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

Education is an important determinant of the length of parental leave. Using the couples sample, 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 occurred over the entire range of the distribution,

[23] When considering the full population aged 20 to 60, this conditional coverage must be scaled by the probability of being employed during the reference month. Individuals not observed in the survey month include students, the unemployed, individuals on sickness or disability benefits, individuals on parental leave, and others without active employment. A reasonable estimate of monthly employment rates among individuals aged 20 to 60 lies in the range of 70 to 80 percent. Using the midpoint value of 75 percent implies an expected overall coverage rate of approximately 50 percent (0.75×0.675).

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. This fraction decreased significantly after the reform. Women's parental leave distributions show a significant reduction between these two periods in long-duration leave, though the absolute change is smaller than that of men's.

Figure 2: Parental Leave Distributions by Education and Sex



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

Education and gender are both important correlates of gender role attitudes. Classifying individuals into university (u) versus non-university (n) categories according to whether they have at least three years of university studies versus fewer, we see significant differences between these groups both pre and post-reform. 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.[24]

Given the importance of education for both lengths of parental leave and gender-role attitudes, we use the university/non-university partition of education to classify couples. This necessarily leaves us with four household types, denoted by nn, nu, un, and uu, where the first entry indicates the father's education and the second the mother's.[25] As can be seen in Table 1, in the pre-period nn couples have men taking the shortest leave and women the longest leave whereas men in uu couples take the longest leave and women the shortest leaves. For mixed education couples, what matters is the education of women in the sense that university women take the shorter leave (and their n partners take longer), whereas non-university women take longer leaves (and their u partners take shorter). The rankings by household type are the same in the post period. Table A1 presents summary statistics for both the wage sample and the couples sample.

Table 1: Mean Months of Parental Leave

		Pre-Period	**Post-Period**
Men	nn	1.99	2.97
	nu	2.62	3.83
	un	2.31	3.45
	uu	3.25	4.54
Women	nn	13.67	12.88
	nu	12.79	12.16
	un	13.27	12.34
	uu	12.02	11.30

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

4 A Simple Model of Parental Leave

As noted previously, the objective of the daddy month reform was to increase the share of parental leave taken by fathers. There are several reasons this could happen. First, there is

[24]Unfortunately, we do not have any questions regarding parental leave itself. Source: simple averages calculated from ISSP (The International Social Survey Programme), 2002, 2012. See ISSP Research https://doi.org/10.4232/1.12661).

[25]Table A2 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 do not affect our analysis as all the moments used for estimation are *by household type*.

a direct incentive effect. The daddy month is an extra month which can be taken for close to full pay. Furthermore, 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 will be shown in Section 5.2, 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's enjoyment of parental leave may depend on how other men behave. As more men respond to the direct financial incentives associated with daddy months, other fathers may find it more rewarding to take additional parental leave themselves.[26] This would also affect mothers' behavior as their desire to take more months of low-paid parental leave might change if their partners are now taking more leave. Fewer women taking lengthy leaves would in turn affect an individual woman's desire to take lengthy leaves herself.

4.1 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 parental leave 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

[26] 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."

preferences can be modeled as:

$$V_{ij} = U(\tilde{y}_{ij}) + B \cdot U(y_{ij}) + \gamma_m^{ij} \cdot h_m^e(t_m) + \gamma_f^{ij} \cdot h_f^e(t_f) \tag{1}$$

where $B > 0$ and i, j denotes the household type (i.e., the education i of the man and j of the woman, $i, j \in \{n, u\}$).[27] 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 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. We assume that utility over income is the same for all individuals.

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 $h(t)$.[28] 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 $\mathcal{N}(\mu_\gamma^{g,ij}, (\sigma_\gamma^{g,e})^2)$ whose mean evolves according to the following law of motion:

$$\mu_s^{g,ij} = a^{g,e} + b^{g,e} \cdot \bar{t}_{s-1}^{g,ij} \tag{2}$$

where $\bar{t}_{s-1}^{g,ij}$ denotes the mean parental leave taken in the previous period, $s-1$, by individuals

[27]In order to simplify notation, equation 1 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.

[28]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.

of the same gender g and household type ij. The standard deviation, $\sigma_{\gamma}^{g,e}$, which also depends on gender and education, is assumed to stay constant over time. Note that the parameters in equation 2, $a^{g,e}$ and $b^{g,e}$ are allowed to depend on gender and education as suggested by the poll evidence. 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 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.[29]

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 2, 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.[30] 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.

The last assumption we make is the usual one for dynamic models: we assume that the economy is in steady state in the pre-reform and the post-reform periods (i.e., μ is constant) and we estimate the model to fit moments for both steady states.

[29] Section 10.2 in the Appendix provides microfoundations for the distribution of γ.

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

4.2 The Maximization Problem

The couple chooses their parental leave durations, t_m and t_f, to maximize their welfare in equation 1 subject to three constraints. First, 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. In our estimation, we set $\tau = 17$ 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 $g \in \{m, f\}$. Second, let T be the maximum possible number of highly paid months. If the couple jointly takes more than T months of leave and each partner takes at least D months (the daddy months), then only the first T months are paid, and any additional month is paid zero.[31] Third, if a parent j takes fewer than D months, then the aggregate parental leave constraint at which pay becomes 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.[32] 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-university woman in the pre-period. The corresponding numbers for the post-period are 45% and 3.6%, respectively.[33] 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.

[31] 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.

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

[33] 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 A1). We do not have data on the costs of private child care. We assume that the cost of private daycare equals the full (unsubsidized) cost of 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.

Note that we have described only how parental leave choices affect income during the 17 month period. Future income, as discussed in Section 5.2, is also a function of these choices as described below.

5 Estimation

This section describes how we parameterize and estimate the model including the wage penalties. 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 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.

Table 2: Description of Parameters

Categories	Parameters	Symbol	Value
External parameters	Mean of log wage draws by gender & hh type	$\mu_w^{g,ij}$	
	SD of log wage draws by gender & hh type	$\sigma_w^{g,ij}$	see Table A3
	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 (private)	C_r	
	Pre-reform period		0.50
	Post-reform period		0.45
	Monthly cost of childcare over age of 1 (public)	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. C_r indicates the private cost and C_u the public cost of childcare. See section 4.2 for more details. The γ distribution is assumed to be normal. See text for definitions of all variables.

5.1 Wage Parameters

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.[34] 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.[35]

5.2 Wage Penalties and Future Wages

Estimating Wage Penalties

Next, we estimate the wage penalties associated with different durations of parental leave.[36] These wage penalties are potentially an important input into the couple's leave decisions. There is now a large literature that examines the effect of parenthood on gender wage and earnings gaps (see Angelov, Johansson and Lindahl (2016) and Kleven, Landais and Søgaard (2019)). Our focus here is different: we are interested in how parental leave duration affects

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

[35] 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.

[36] Parental leave can also be costly for employers, e.g., the costs associated with hiring replacement workers. Ginja, Karimi and Xiao (2023) uses a 1989 reform that allowed workers to postpone their return to work by an additional 3 months. The authors find that this reform significantly affected firms' labor costs whereas they find that the 2002 reform, which mostly affected fathers, did not.

the within-gender and education distribution of wage growth.[37,38]

To estimate wage penalties, we use the full wage sample described earlier and examine wages both three years after and one year before the birth of the first child. The underlying assumption is that an individual's wage penalty is driven by their gender and education rather than household composition. All wages are expressed in 2014 SEK. Wage equations are specified separately for each gender-education group and for each period (pre and post-reform).

Accordingly, we have the following wage equations for these two periods:

$$\log(w_{i,t+3}^{g,e}) = \alpha_{t+3}^{g,e} + \beta_1^{g,e} plcat_i + \beta_{2,t+3}^{g,e} agecat_i + \chi_i^{g,e} + \lambda_{t+3}^{g,e} + \epsilon_{i,t+3}^{g,e}$$
$$\log(w_{i,t-1}^{g,e}) = \alpha_{t-1}^{g,e} + \beta_{2,t-1}^{g,e} agecat_i + \chi_i^{g,e} + \lambda_{t-1}^{g,e} + \epsilon_{i,t-1}^{g,e}$$

where $\log(w_{i,t}^{g,e})$ denotes the log wage of individual i of gender $g \in \{m, f\}$ and education $e \in \{n, u\}$ in year t. $\alpha_t^{g,e}$ is a time-varying intercept. The variable $plcat_i$ is a vector of categorical indicators (defined below) for the number of months of parental leave taken by individual i within 16 months following childbirth (with the associated wage penalties represented by the vector $\beta_1^{g,e}$). This variable is excluded from the second equation, as parental leave cannot be taken in year $t-1$. The term $agecat_i$ denotes a vector of age-at-childbirth dummies, one for each age, with time-varying coefficients in the vector $\beta_{2,t}^{g,e}$. The specification includes individual fixed effects ($\chi_i^{g,e}$), as well as sample year fixed effects $\lambda_t^{g,e}$, which also capture variation across childbirth cohorts which are not separately identified. $\epsilon_{i,t}^{g,e}$ denotes the error term.

A potential concern is selection into different durations of parental leave. For instance, individuals with weaker labor market performance prior to childbirth may be more likely to take longer leave and to continue to earn lower wages thereafter. Estimating the model in first differences helps mitigate this concern by removing time-invariant individual characteristics, whether observed or unobserved. However, we do not interpret the resulting

[37] Albrecht et al. (1999) is an early article that estimates the effect of time off work (including, but not limited to, parental leave) on future wages.

[38] It is worth noting that these wage penalties are not to be confused with motherhood penalties for women. The wage penalty estimates the difference *across mothers* in their wage growth after three years as a function of the *length of their parental leave* rather than by how much their wage changes as a result of becoming a mother.

estimates causally, as selection on time-varying unobservables remains a potential source of bias.

Subtracting the wage equation at $t-1$ from that at $t+3$, yields:

$$\log(w_{i,t+3}^{g,e}) - \log(w_{i,t-1}^{g,e}) = \alpha_0^{g,e} + \beta_1^{g,e} plcat_i + \beta_2^{*g,e} agecat_i + \phi^{g,e} + \epsilon_i^* \tag{3}$$

which provides estimates of the wage penalties, $\beta_1^{g,e}$, three years after childbirth. Here, $\beta_2^{*g,e} = \beta_{2,t+3}^{g,e} - \beta_{2,t-1}^{g,e}$ captures the change in the returns to age across the two periods, and $\phi^{g,e} = \lambda_{t+3}^{g,e} - \lambda_{t-1}^{g,e}$ captures the change in year (or cohort) fixed effects. ϵ_i^* denotes the change in the error term over time. Finally, $\alpha_0^{g,e}$ represents the new intercept and captures the wage difference for the reference category. 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 Figure 5.

Table 3 reports these wage penalties averaged within intervals of months over which the coefficients are similar (but not all individually significant). The table shows several interesting patterns. For non-university men, there are penalties associated with taking any amount of parental leave. The penalty is smaller in the post-period than in the pre-period and, in both cases, rises with parental leave uptake. In both the pre- and post-periods, university men can take several months of leave without incurring any penalty, but there is a substantial penalty associated with taking relatively long leave.

Non-university women in the pre-period suffer no wage penalties for taking up to 11 months of leave. Those who take more than 11 months suffer relatively small penalties. Penalties for these women are even smaller in the post-period. Finally, among university women in the pre-period, there are penalties associated with taking 11 months of leave or more. The penalties are substantially higher for long leaves, i.e., 17 months. 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.

In sum, on average, parental leave penalties for both men and women decreased between the pre- and post-periods. 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 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 the public sector or to 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, 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.[39] Lastly, a smaller percentage of women switched employers at t+3: 36% in the pre-reform period versus 26% in the post-reform period.[40] 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.[41] 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 (e.g. those of Section 6) are very similar.[42]

A potential concern is that individuals may have different wage trends before $t-1$ if they were anticipating having a child and that the wage penalties are simply reflecting this. To allay this concern, we have re-estimated the wage penalties under two alternative specifications: one controlling for wages at $t-2$ and another for wages at $t-2$ and $t-3$. As can be seen in Appendix Figure 5, these alternative specifications yield very similar results.

Future Wages

Given each draw of $\log(w^{g,e}_{i,t-1})$, we construct future log wages using equation 3, the couple's parental-leave choices, and the wage penalties in Table 3. Letting $\widehat{\ }$ denote estimates, we

[39] Here we denote "full time" as working more than 75% of standard "full time."

[40] 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 these data with the individual-level data used in our analyses.

[41] 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.

[42] Results available from the authors.

Table 3: 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.

calculate:

$$\widehat{log(w_{i,t+3})} = \widehat{\alpha}_0^{g,e} + log(w_{i,t-1}) + \widehat{\beta}_1^{g,e} plcat_i + \widehat{\beta}_2^{*g,e} a\bar{g}e^{g,e} + \widehat{\phi}^{g,e} \tag{4}$$

where we evaluate the year fixed effects at the mean and the age effect at the median age for each gender-education group. Appendix Table A4 reports the relevant values of $\widehat{\alpha}_0^{g,e}$, $\widehat{\beta}_2^{*g,e}$, and $\widehat{\phi}^{g,e}$ by gender and education for the pre and post periods.

Lastly, we exponentiate the predicted future log wages, normalize as before, 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 3 and of any childcare costs 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.

5.3 Other Functional Forms

We assume that utility over income (current and future) and over the length of time spent with a child during parental leave are given by CRRA preferences:

$$u(y) = \frac{y^{1-\eta}}{1-\eta}, \quad h_{g,e}(t) = \frac{t^{1-\zeta_{g,e}}}{1-\zeta_{g,e}} \tag{5}$$

with $\eta, \zeta \geq 0$. The constant (B) multiplying the utility function over future household income in equation 1 is the same for all households as is the CRRA utility function. For utility over length of parental leave, the curvature is allowed to be a function of both gender and education. This implies, for example, that a non-university man has the same curvature over parental leave time independently of whether his partner is a university or non-university woman. The parameters governing the distribution from which individual γ values are drawn (σ, a, and b) also depend only on gender and education, but the evolution of the mean over time will depend on peers' behavior and hence on an individual's household type. This allows the preferences over leave to differ, endogenously for, say, a university man with a university partner relative to a university man with a non-university partner.

5.4 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 the steady states of 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 5 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 male to female household wage ratio and the within-household share of leave taken by men (2x4 moments), totaling 56 moments.[43] In particular, it is important for the model to capture the fact that, despite economic 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.[44]

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

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

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

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 of very low paid leave which we set to zero). Using the specification of household utility of Equation 1, 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 10.5. 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 4 reports the values of the estimated parameters along with their standard errors.[45]

5.5 Estimation Results and Parameter Interpretation

As can be seen in Table 5, the model does a very good job matching the data.[46] 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.[47] 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

[45] 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.

[46] The mean over the weighted sum of squared errors is 0.013.

[47] 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.

Table 4: 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.

the share of leave taken by men.

It is useful to provide an interpretation of some of the parameters reported in Table 4. First, note that η – the parameter governing the curvature of the utility function over income, common to all – is slightly greater than one (which would give log utility). The parameter B, which weights the utility of income three years later, equals 2.65, indicating that household income three years later – 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.[48] 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 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

[48]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.

Table 5: 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.

gender, but greater for men than for women. The model requires this to accommodate the fact that the standard deviation of men's parental leave months is large relative to its mean in the data, whereas the standard deviation of women's parental leave months is small relative to its mean (see Table 5).

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}.[49] 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:$U(y^{ij}) - U(y^{ij}(1-x)) = \mu_\gamma^{g,ij} \cdot (h_{g,e}(7) - h_{g,e}(6))$ 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 6.[50]

Not surprisingly, as can be seen in the second row of Table 6, 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 increases 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 6. 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.

[49] 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.

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

Table 6: 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.

5.6 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).

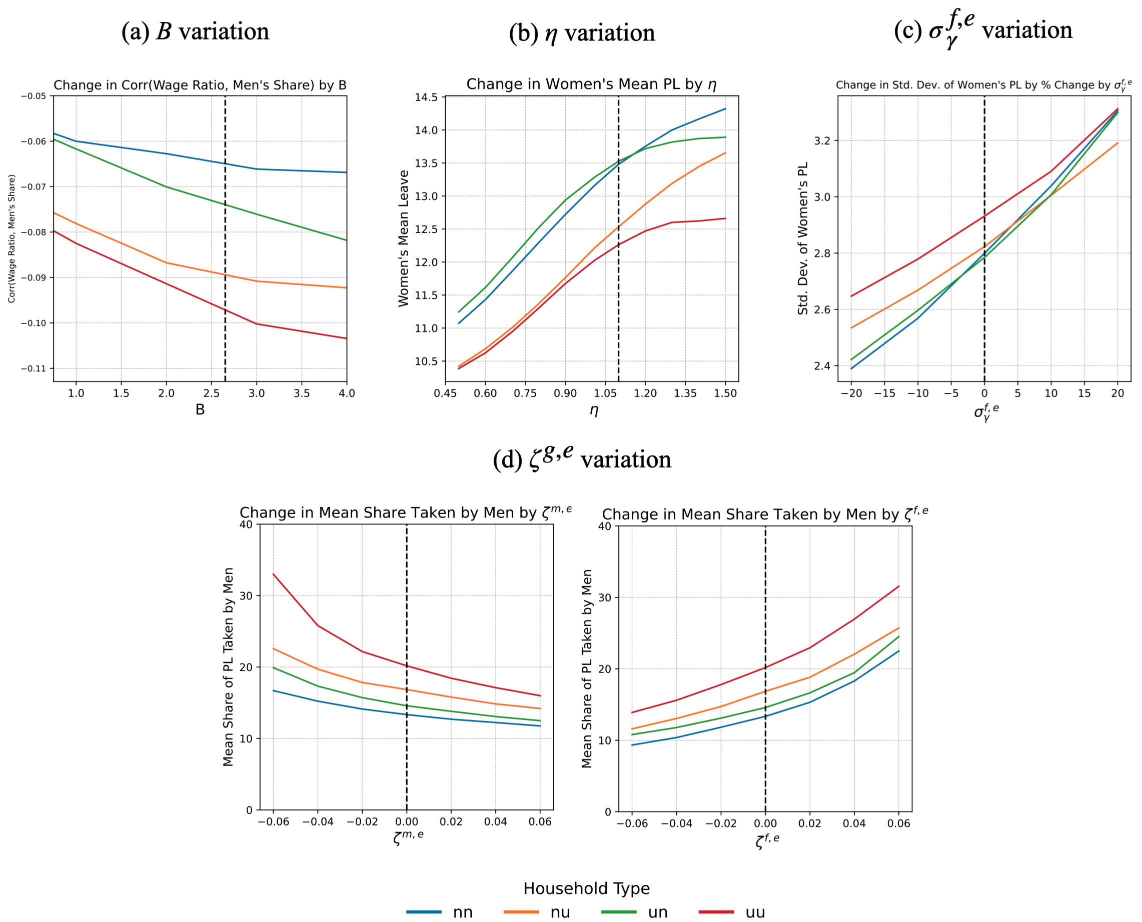
Figure 3 illustrates how large variations in $B, \eta, \sigma_\gamma^{f,e}$, and $\zeta^{g,e}$ affect given equilibrium outcomes in the pre-reform period, keeping all other parameters fixed. As can be seen in Figure 3(a), higher values of B are associated with higher (in absolute value) correlations between the household gender wage ratio and the man's share of leave for all household types. The intuition is that higher values of B increase the importance of *future* income and the opportunity cost of parental leave, strengthening the conventional economic relationship between relative household wages and who foregoes income to stay home with a baby since men tend to earn more and face higher wage penalties. Turning next to increases in η, as shown in Figure 3(b), the mean leave of women increases. 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, Figure 3(c) graphs the effect of changes in the variance of 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 increase the standard deviation of women's parental leave since tastes for the latter are now more dispersed. Lastly, Figure 3(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.[51] 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 men's share as women want to spend less time on leave.

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

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

Figure 3: 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.

taste distribution for leave implies that a significant fraction of men do not wish to take leave and the curvature of the utility of time with children will then have a large effect on the proportion that chooses a corner solution of zero as well as mean leaves.

6 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. Note that it is only when we allow norms to evolve that there are any dynamics; in that case we solve for the steady state.

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 while it increased the number of months a man would "throw away" by not taking parental leave. The consequences of the reform alone, i.e., keeping all income-related variables at their pre-reform levels and not allowing culture to change, are shown in column 2 of Table 7, 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 ranging from .11 to .20 of a month. It generates from 9% (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 and wage penalties fell. 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 generate 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%. Appendix Table A5 examines how each income-related change separately affects parental leave. For men the most important dimension is the decrease in wage penalties whereas for women they all play a minor role.

It is important to note that, for both this experiment and the previous one, the percentage of men who take 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 means of the γ distributions to change endogenously in response to the reform and solving for the new steady state. As shown in column 4, 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 men. For women, the additional contribution is also very large, especially for women with a non-university partner as it accounts for 45% (nn) and 53% (nu) of their total leave decrease.

Table 7: 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 3 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.

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 university 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.

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. 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.

7 Empirical Validity and Income vs Culture

We next turn to exploring two facets of our results. First, we examine how the model-derived magnitudes of the effects of only the reform 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.

The Short-Run Effect of the Reform: Contrast with an RD Analysis

In the prior section, we isolated the effect of the reform that came solely from changing individuals' opportunity sets. To do this, we imposed the reform but kept constant all income-related variables as well as parental leave preferences at their pre-reform levels. To examine the empirical validity of our model, we contrast the short-run effects predicted by the model with those obtained from a regression discontinuity analysis and, as we show below, obtain similar results.

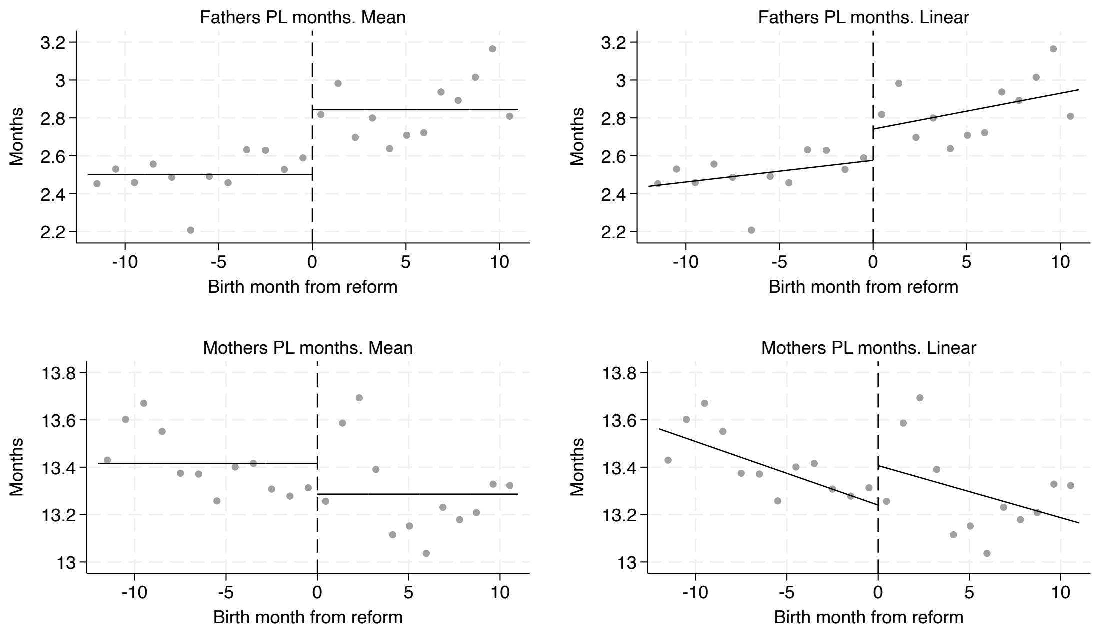
In particular, we exploit the discontinuity introduced by the 2002 reform to estimate its impact on parental leave uptake using a sharp regression discontinuity (RD) design. To be consistent with our model estimation, we measure parental leave as the number of leave months taken within the first 16 months following childbirth. To ensure a large enough sample, we use all individuals with wages in $t-1$.

We use the MSE-optimal bandwidth procedure described in Calonico, Cattaneo and Farrell (2020) to estimate the regression discontinuity effects separately for fathers and mothers. The RD samples consist of 23,626 fathers and 30,954 mothers. Applying the bandwidth selection procedure, which trades off bias and variance, results in smaller effective sample sizes that differ across specifications, as reported in the RD regression table below. Figure 4 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 0.28 months in the mean and linear specifications. In contrast, for mothers there is no clear evidence of a discontinuity associated with the 2002 reform.[52] The RD estimates are reported in Table 8.

To contrast the RD results with our model generated ones, we 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 environment in 2001.

[52] 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 increased their leave by an average of six to seven days, based on a 24-month follow-up after birth. Focusing on fathers, the coefficient of 0.28 implies an extra 8.4 days of leave. Our estimates are smaller than those of Avdic and Karimi (2018), likely reflecting our shorter, 16-month post-birth observation window. On the other hand, our estimates are somewhat larger than those of Duvander and Johansson (2012). This likely reflects the fact that we use first births rather than all births.

Figure 4: 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 full sample of births in 2001 and 2002, restricted to individuals with observed wages in the year prior to childbirth.

Table 8: 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	0.279	0.282	0.080	-0.071
Robust_SE	0.157	0.238	0.156	0.307
Robust_z	1.781	1.184	0.511	-0.231
EffN_left	827	2624	2147	2147
EffN_right	1905	4095	3732	3732
EffN_total	2732	6719	5879	5879
Polynomial	Mean	Linear	Mean	Linear
Bandwidth	MSE-optimal	MSE-optimal	MSE-optimal	MSE-optimal

Note: The table reports regression discontinuity (RD) estimates of parental leave usage within 16 months following childbirth, estimated separately for fathers and mothers with wages in $t-1$. Estimates are obtained using a triangular kernel and MSE-optimal bandwidths on each side of the cutoff. Both local constant ($p=0$) and local linear ($p=1$) specifications are reported. Robust bias-corrected standard errors are used to construct the reported z-statistics. Effective sample sizes (EffN_left, EffN_right and EffN_total) refer to kernel-weighted observations within the selected bandwidths and may differ from the total number of observations. Because the running variable is discrete, the MSE-optimal bandwidth often maps to the same set of monthly support points across specifications and polynomial orders (see Columns 3 and 4).

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 (culture) at their 2001 pre-reform values. This exercise is analogous to the "Reform Only" column in Table 7, 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 A6.

Table 9 reports the results of this exercise. As shown in the table, men witnessed an increase of 0.21 months whereas women saw a negligible decrease of 0.03 months of parental leave. These results are fairly close to the estimates obtained from our RD, lending support to our model's ability to explain the data.

Table 9: RD Simulation Results

	Men	Women
Pre-Reform, 2001	2.32	13.18
Post-Reform, 2002	2.53	13.13
Difference	0.21	-0.05

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 A6. The "Difference" row indicates the change in leave uptake (post minus pre).

A Counterfactual Correlation: Can Income be the Dominant Force?

A legitimate question is whether the model could allow culture 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 A7 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.[53] In that counterfactual estimation, the reform on its own has only a small effect on parental leave. The additional contribution of income-related changes over the reform is large, ranging from 49 to 89%. Furthermore, while the percentage of men taking zero leave changes very little for non-university men, it significantly decreases 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 3 and 5, 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 alone had certain key moments been different from those in the data.

8 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.[54] 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.[55] In Sweden, the reform that we study increased men's share of leave, but still 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 the effect on gender equality of i) equalizing the male and female wage distributions and ii) alternative childcare and parental leave policies.

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,

[53] This counterfactual estimation has a significantly larger mean of weighted sum of squared errors of 0.031 as opposed to 0.013.

[54] 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.

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

wage penalties, and childcare costs as in the post-reform environment. While these may well respond to any reform, we do not have a model of wage determination or signaling that allows us to endogenize the latter. 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.

8.1 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' wage in the year prior to the birth of their first child.[56] We first examine the role of the gender wage gap on the 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 the wage distribution of university men in un households. The household correlation of wages and the level of wage penalties by education and gender are kept unchanged from their values in the post-reform period.

Columns 2 and 3 of Table 10 report the results of this exercise. For ease of comparison, column 1 reproduces the simulated results of the two-daddy months reform. Not surprisingly, the effect 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), changing men's share of parental leave by around 0.5 to 1 percentage point. Once preferences are allowed to change endogenously (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 powerful force for promoting gender equality in parental leave uptake.

[56] Calculated from the couples sample, year t-1 of the post-reform period.

Table 10: 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.

8.2 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.[57] 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.[58] 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 10, providing inexpensive childcare during the child's first year has almost no effect on parental leave. Women's mean months of leave 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 given the lack of behavioral responses to this policy, which explains why both columns 4 and 5 show virtually 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.

Next, we examine the 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 11. Columns 2 and 4 report the moments resulting from the endowment and equal-sharing policies, respectively, assuming that parental leave 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

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

[58] 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.

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 of around 2 to 2.5 months. 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.[59] 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.[60] 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 sharing 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 6.5 - 6.8 months of parental leave whereas university men take around 6.9 - 7.4 months. The percentage of men who take zero months of leave is under 2% for all household types. Under both policies, women take similar amounts of leave, though slightly lower under equal sharing. Thus, although incentives with invariant preferences gave rise

[59]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.

[60]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.

Table 11: 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.

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 11 shows the mean values of the $\mu_{\gamma}^{g,ij}$ for each of the five columns. The steady state values under the endowment and equal share policies (columns 3 and 5) 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 in the new steady states are now higher than women's. This does not imply, however, that men prefer to spend more time with children than women do. The curvature of the utility function over time with one's child also plays an important role. As can be seen in Table 4, 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.[61]

9 Conclusion

Even in countries in which gender equality is more advanced, care responsibilities continue to follow traditional gender roles. 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, it may have important intergenerational consequences that further change attitudes and economic outcomes.[62]

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 mediating mechanisms, we develop

[61] In our model of cultural change, we assumed that the extensive part of preferences – the $h(t)$ functions – are invariant. One might also imagine that this would change over time, bringing men's and women's preferences closer together.

[62] As shown by Kotsadam and Finseraas (2013) and Farré et al. (2023) for Norway and Spain, respectively.

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 is slightly higher. 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.

Our counterfactual policy analysis has implications for other countries. It cautions that simply making childcare more affordable is unlikely to generate large changes in parental behavior, at least not in countries where women already have a strong attachment to the labor force. Instead, earmarking parental leave so that parents face stronger incentives – both economically and, over time, socially – to share leave more equally is likely needed to produce greater equality.

10 Appendix

10.1 Summary Statistics

Table A1: 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.

Table A1 presents summary statistics for the wage sample and couples sample discussed in Section 3 above. Table A2 gives the distribution of household types in the couples sample by period.

Table A2: 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.

10.2 Some Microfoundations for the Normal Distribution of γ

We can model individual taste for parental leave, γ, as

$$\gamma_{i,s}^{g,ij} = \mu_s^{g,ij} + \varepsilon_i^{g,e}$$

where $\mu_s^{g,ij}$ is a common cultural component that moves with last period's mean and $\varepsilon_i^{g,e}$ is idiosyncratic heterogeneity of mean zero and constant variance that is time invariant. For each individual, the idiosyncratic term is the sum of several small, zero-mean, roughly independent influences, $\zeta_i^{(k)}$, i.e., $\varepsilon_i^{g,e} = \sum_{k=1}^{K} \zeta_i^{(k)}$, reflecting factors such as who their supervisor is at work, partner logistics, health, etc. with no single factor being dominant. In that case, we can use the within-person Central Limit Theorem which implies $\varepsilon_i^{g,e} \overset{approx}{\sim} \mathcal{N}\big(0, (\sigma_\gamma^{g,e})^2\big)$. Thus,

$$\gamma_{i,s}^{g,ij} | \mu_s^{g,ij} \sim N(\mu_s^{g,ij}, \sigma_\gamma^{g,e})$$

10.3 The Couple's Maximization Problem

Using the notation in Section 4.2, we are now set to express household income for both the pre- and post-reform periods during our $\tau = 17$ month observation window. 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, although 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 taken by the parent whose wage is greater and by $\underline{t}$ the parental leave taken by the parent whose wage is lower. Lastly, we denote by t_d, w_d the parental leave 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 and wage of the other spouse.[63] Depending on the total length of parental leave 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 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 $D = 2$.[64]

A second case is for the couple to jointly take more than T months, i.e., they take some

[63] 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 described below.

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

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 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 Wages

Wage Penalties

The wage penalties in Table 3 are constructed by averaging coefficients over contiguous intervals in which the point estimates are similar, as described in the text. The graphs in

Figure 5 below show the estimated coefficients and 95% confidence intervals.

The parameters of the log normal distribution of wages are taken directly from the data. Table A3 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 Monthly Wage $(t-1)$ in Table A1 for the couples sample.)

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

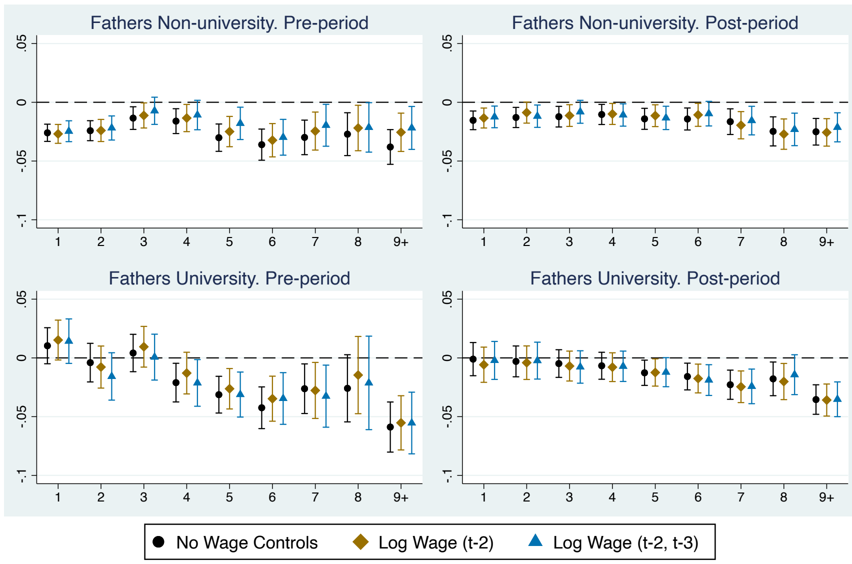
Note: These are the means and variances of real (2014 SEK) log wages for each gender-household type along with the correlation of partners' log wages. All variances on this table are multiplied by 10. We used real log wages from the year prior to the first birth for both the pre-reform period and the 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

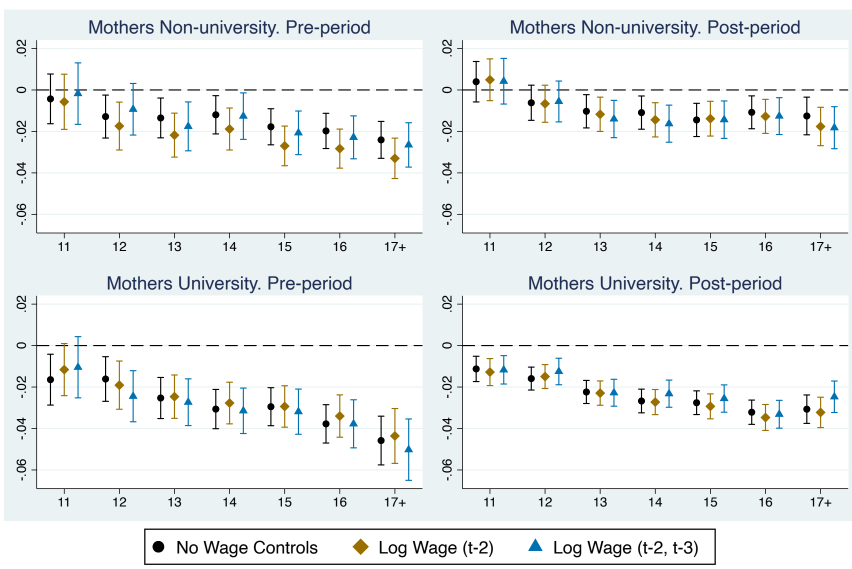
Table A4 shows the estimates used in constructing future wages by gender and education.

Figure 5: Wage Penalty Estimates

(a) Men



(b) Women



Note: Coefficients from the estimation of equation 3 for our baseline are in circles (black - no wage controls). Two alternative specifications – one controlling for wages 2 years before birth of child (gold diamonds) and the next controlling for both 2 and 3 years prior to birth (blue triangles) – are included for comparison. 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. 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.

Table A4: Wage Growth Parameters

	Men							
	Non-Univ				Univ			
Period	$\widehat{\alpha}_0^{g,e}$	$\widehat{\phi}^{g,e}$	$\widehat{\beta}_2^{*g,e}$	$\bar{\text{age}}^{g,e}$	$\widehat{\alpha}_0^{g,e}$	$\widehat{\phi}^{g,e}$	$\widehat{\beta}_2^{*g,e}$	$\bar{\text{age}}^{g,e}$
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	$\widehat{\alpha}_0^{g,e}$	$\widehat{\phi}^{g,e}$	$\widehat{\beta}_2^{*g,e}$	$\bar{\text{age}}^{g,e}$	$\widehat{\alpha}_0^{g,e}$	$\widehat{\phi}^{g,e}$	$\widehat{\beta}_2^{*g,e}$	$\bar{\text{age}}^{g,e}$
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: $\widehat{\alpha}_0^{g,e}$ is the constant term in the wage penalty regression, as shown in equation 3. For each gender-education group, $\widehat{\phi}^{g,e}$ is the average year fixed effect, $\widehat{\beta}_2^{*g,e}$ is the coefficient on the median age dummy, and $\bar{\text{age}}^{g,e}$ is the median age, as shown in equation 4.

10.5 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.[65]

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.[66] This iterative process continues until we find a parameter vector that minimizes our objective function.

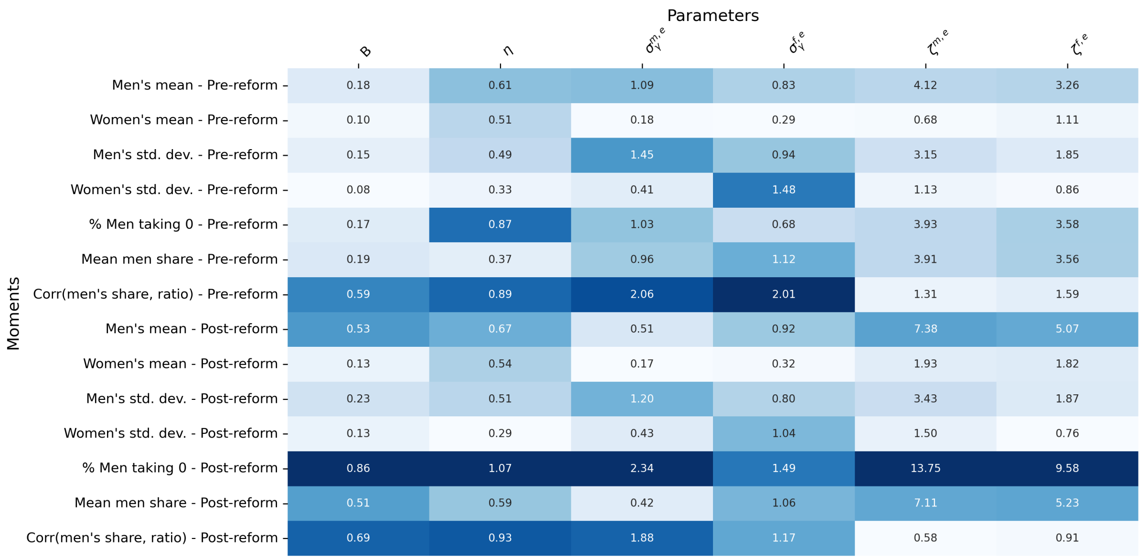
10.6 Derivative Analysis

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 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.

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

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

Figure 6: Derivative Analysis



Note: We perturb each structural parameter individually by 1% while holding all other parameters fixed and measure the resulting percentage change in each simulated moment. Each column (parameter) 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.

10.7 A Finer Decomposition of Income-Related Changes

Table A5 examines the contribution of the reform with different aspects of income-related changes between the two periods: the wage distribution, childcare costs, and the wage penalties.

10.8 Regression Discontinuity

The model simulation of the RD use wage distributions by household type from 2001 (pre-reform) and weights from 2001 and 2002 used to aggregate the model's household-type–specific means into single gender-specific means. These are reported in Table A6.

10.9 Counterfactual Correlation

Table A7 gives 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

Table A5: 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 3. 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.

Table A6: RD Parameters for Model Simulation

Panel A: Pre-Reform 2001
Log Wage Parameters

	nn	nu	un	uu
Men				
$\mu_w^{m,ij}$	9.89	9.98	10.08	10.09
$\sigma_w^{m,ij}$	0.25	0.28	0.33	0.32
Women				
$\mu_w^{f,ij}$	9.72	9.83	9.86	9.95
$\sigma_w^{f,ij}$	0.20	0.22	0.26	0.26

Panel B: RD Sample Observations

	nn	nu	un	uu
Men				
Pre-Reform, 2001	6597	1847	960	2034
Post-Reform, 2002	6481	2100	1147	2460
Women				
Pre-Reform, 2001	8500	3137	977	2391
Post-Reform, 2002	8313	3550	1095	2991

Panel C: Weights

	nn	nu	un	uu
Men				
Pre-Reform, 2001	0.58	0.16	0.08	0.18
Post-Reform, 2002	0.53	0.17	0.09	0.20
Women				
Pre-Reform, 2001	0.57	0.21	0.07	0.16
Post-Reform, 2002	0.52	0.22	0.07	0.19

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.

of -0.60. As can be seen, the model does a bad job in matching several moments related to the length of women's parental leave (e.g., obtaining close to 9.5 months in the pre-reform and post-reform periods rather than the 12-13.5 pre-reform months and 11-13 post-reform months given in the data) and is forced to leave the correlation closer to -0.4 to provide a lower sum of squared errors than it would achieve by hitting the targeted moment.[67]

[67] The mean over the weighted sum of squared errors is considerably higher, 0.031 as contrasted with 0.013.

Table A7: 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 3, 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.

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.

Calonico, Sebastian, Matias D Cattaneo and Max H Farrell. 2020. "Optimal bandwidth choice for robust bias-corrected inference in regression discontinuity designs." *The Econometrics Journal* 23(2):192–210.

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.

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. Working Paper 34077 National Bureau of Economic Research. forthcoming in the Handbook of Culture and Economic Behavior, edited by Enke, B and P. Giuliano and N. Nunn and L. Wantchekon, Elsevier.

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

Ginja, Rita, Arizo Karimi and Pengpeng Xiao. 2023. "Employer Responses to Family Leave Programs." *American Economic Journal: Applied Economics* 15(1):107–135.

Gonzalez, L and H. Zoabi. 2023. Does Paternity Leave Promote Gender Equalty within Households. Working Paper 9430 cesifo.

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.