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Abstract

In this paper we use a parental leave reform directed towards fathers to identify the causal effects of paternity leave on children's and parents' outcomes. We document that paternity leave causes fathers to become more important for children's cognitive skills. School performance at age 16 increases for children whose father is relatively higher educated than the mother. We find no evidence that fathers' earnings and work hours are affected by paternity leave. Contrary to expectation, mothers' labor market outcomes are adversely affected by paternity leave. Our findings do therefore not suggest that paternity leave shifts the gender balance at home in a way that increases mothers' time and/or effort spent at market work.

JEL-Code: J130, J220, J240, I210.

Keywords: parental leave, labor supply, child development.

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1 Introduction

Paternity leave is often discussed as a policy measure to encourage greater gender equality both in the family and in the labor market. Politicians and policymakers in Northern Europe are strong believers that paternity leave strengthens women's position in the labor market, reduce the gender wage gap and give children a chance to bond with their fathers - and vice versa.[1]

Wishing to alter the traditional household specialization, politicians like to provide incentives to increase men's involvement in the home. Even a few weeks of paternity leave, the argument goes, may result in substantial changes.[2] Thus recently Finland, Iceland, Norway and Sweden have all reserved a share of the parental leave for fathers. Similar proposals are also popular and highly debated in other European countries.

In this paper we challenge the popular view. Our paper investigates how paternity leave impacts a broad range of outcomes, using Norwegian register data. To handle the selection problem we use the introduction of the paternal quota in Norway on April 1, 1993, to evaluate the causal effects of paternity leave on children and parents. We find that *children's school performance* benefits from paternity leave in families where the father is relatively higher educated than the mother. Consistent with this finding, *fathers' earnings and working hours* seem to be negatively affected by paternity leave, but the effects are not statistically significant. Contrary to expectation, there are strong and statistically significant negative effects on *women's labor market outcomes* of their spouse taking paternity leave. Furthermore, paternity leave has no significant effect on a set of *family outcomes* such as fertility and divorce rates.

Time allocation data from the United States show that fathers of sons spend more time with their children than fathers of daughters (see survey by Lundberg (2005)). Fathers

[1] These views are articulated in a series of white papers, cp. 'Likestilling for Likelønn' (Stortingsmelding nr. 6 (2010-2011)) and 'Reformerad Föräldraförsäkring - Kärlek, Omvårdnad, Trygghet' (SOU 2005:73).

[2] "To strengthen the father's role in his child's life, it is important for him to participate in childcare during the child's first year. A portion of the parental leave period should therefore be reserved for the father". (The Norwegian Government's Long term program for 1990-1993 (Stortingsmelding nr. 4), our translation.)

of sons are also found to be more involved in school work than fathers of daughters (Morgan et al. (1988)). In Scandinavia, however, fertility decisions indicate a preference for daughters rather than sons (Andersson et al. (2006)). Furthermore, several studies (surveyed in (Almond and Currie, 2010)) indicate that interventions can have different long-term effects on boys and girls. For instance, girls seem to be more responsive to preschool interventions (Cascio (2009), Havnes and Mogstad (2011)).

Accordingly, one should expect the effects of fathers spending more time with their children to differ according to the child's gender. Indeed, our estimated effect on school performance is driven by an effect on girls' outcomes. For boys the estimates are smaller and statistically insignificant. The difference persists for the other outcomes as well: The fathers and mothers that work and earn less due to paternity leave are the ones that get daughters. The estimated impacts on family outcomes consistently have opposite signs in families who had girls and in families who had boys.

While several papers have investigated how maternity leave (or general parental leave) impacts parent (e.g. Lalive and Zweimüller (2009)) and child outcomes (see Baker and Milligan (2011) for a review), there are few studies that have considered the particular effects of paternity leave. Using paternal quota reforms in Sweden, Ekberg et al. (2005) find no evidence that paternity leave affects the extent to which fathers care for children when they are sick, whereas Johansson (2010) finds no causal effect on mothers' and fathers' earnings. However, the precision in the estimates of both of these studies is very low. Using Norwegian data, Rege and Solli (2010) find a negative effect of paternity leave on fathers' earnings. We are not aware of any previous studies of how children's outcomes are affected by paternity leave.

The rest of the paper is organized as follows: Section 2 discusses the institutional setting and the reform. Section 3 presents our empirical strategy. In Section 4 we describe our data and the various outcome measures that we use. In Section 5 we present the results. Section 6 concludes.

2 Paternity leave in Norway

In Norway, wage compensated parental leave has been extended continually since the 1970s, from 18 weeks leave with full wage compensation in 1977 to 46 weeks in 2009.[3] Although the parental leave scheme offers full (or 80%) wage compensation, eligibility for wage compensated leave is contingent on the *mother* having worked 50% or more during at least six out of the last ten months before the child's birth.[4] This applies to both parents: Norwegian fathers have no independent right to paid parental leave. If the mother works part time (between 50 and 100%), the father's compensation rate is reduced accordingly. In addition, the father must himself have worked at least six out of the last ten months in order to be eligible for wage compensated leave.[5]

Historically, men have taken very little of the leave period that can be freely shared between the parents.[6] This fact was debated in Norway during the 1980's, and various measures that would induce men to take part of the leave were called for. In the autumn of 1992, the labor party government included in their suggestion for the national budget of 1993 a seven-week extension of the (100%) compensated parental leave period, of which four weeks would be reserved for fathers.

The reform was passed in parliament in December 1992. Following implementation on April 1, 1993, four of the 42 weeks of paid leave were reserved for the child's father. Except under special circumstances, families would lose the right to these four weeks unless taken by the father. In addition, the mother had to resume work for the father to be eligible for the paternal quota.[7] All subsequent extensions of the parental leave period

[3]See Appendix Table 15 for a full description.

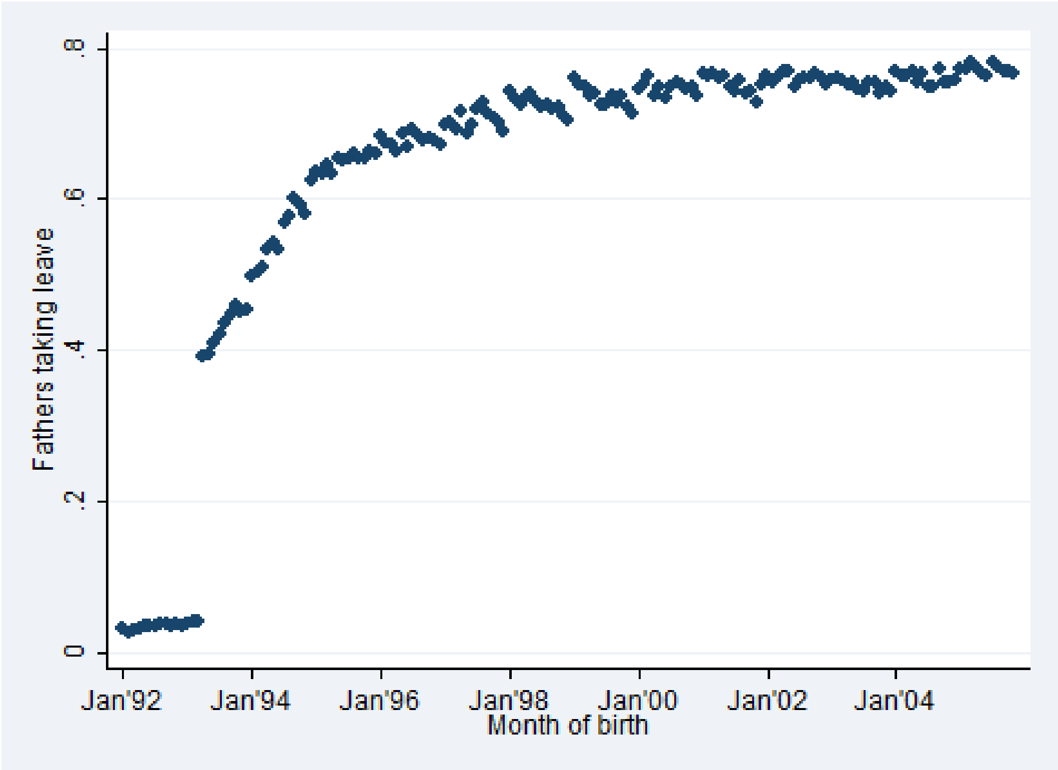
[4]Sick leave from employment, unemployment with right to benefits, and paid parental leave all count as work.

[5]Income compensation also reaches an upper bound of six times the *basic amount* (G) ("Folketrygdens grunnbeløp") of the Norwegian social security system. This amount is adjusted yearly (or more often) in accordance with changes in the general income level. From January 1 2010, G is NOK 72 881 (apprioximately USD 12 500). Until 2008, when self-employed were granted rights to full compensation, the compensation rate for self-employed was at 65% of their income.

[6]The parental leave period can be shared between the mother and father, except for the first six weeks after birth, which have been reserved for the mother. In 1991, women were obliged to start their leave period two weeks before expected delivery. From 1993 this is extended to three weeks.

[7]This requirement was relaxed in July 1994 (Brandth and Øverli (1998))

Figure 1: Share of fathers taking leave in families eligible for parental leave, 1992-2005.



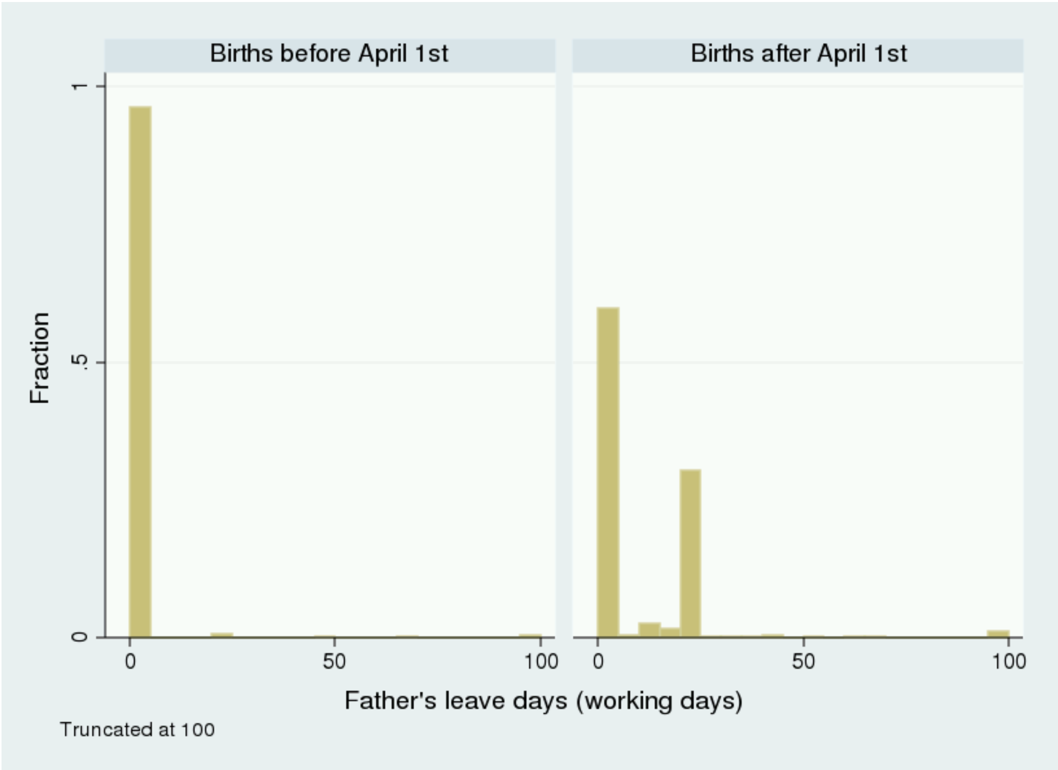
have been put into the paternal quota.

As is evident from Figure 1, the 1993 reform dramatically increased the number of fathers taking paternity leave. In our sample of families eligible for parental leave, the fraction taking paternity leave immediately increased from 4% in March 1993 to 39% in April 1993. The fraction has increased steadily over time, to above 80% of eligible fathers in 2005.

Figure 2 gives the distribution of paternity leave spells in our sample of fathers of children born in a 26 week period surrounding April 1, 1993, in families who are eligible for parental leave. The 40% percent of the fathers eligible for the paternal quota take on average about 25 days of leave (five weeks), with almost three quarters taking exactly the four weeks of the quota.[8] Most fathers only have one spell of paternity leave - if any (90%). On average, their leave period starts when the child is nine months old. Less

[8]Only 10% of leave-taking fathers took more leave than the paternal quota until 1999. The fraction rose to to 18% in 2004.

Figure 2: Fraction of eligible fathers by number of leave days taken (working days).



Note: The sample is fathers of children born in a 26 week period surrounding April 1, 1993, who were eligible for parental leave. A small number of fathers have very many leave days, for ease of exposure the number of leave days have been truncated at 100.

than 5% take leave after the child has turned one year.

3 Identification

Estimating the causal effects of paternity leave on parent, family and child outcomes is complicated by a selection problem. In families where fathers take parental leave, both parents tend to be older, more educated and have higher income than in families where fathers do not take parental leave. These families are also likely to differ along unobservable characteristics as well. We handle the selection problem by exploiting the introduction of the paternal quota at April 1, 1993.

3.1 Reform as exogenous variation

Our empirical strategy is based on the idea that when looking at births closely surrounding April 1, 1993, the paternal quota reform provides quasi-experimental variation in the uptake of paternity leave, since only families with children born after April 1, 1993 are eligible for the paternal quota.

As mentioned in Section 2, not all of these families were actually covered by the reform: Parents must be eligible for wage compensated parental leave for the reform to represent an actual change in incentives. Therefore, in all of our analyses, our samples are restricted to eligible families. How eligibility status is defined and determined will be discussed in Section 3.4.

To eliminate inherent differences between families with children born at different times of the year[9], we rely on a difference-in-differences approach, comparing the difference between the 1993 pre-reform and post-reform cohorts in 1993 to that between corresponding cohorts from 1992.

Families of children born during the same calendar month in the previous year constitute a natural comparison group. Using 1992 as our comparison year is particularly useful to us, as there was a reform extending general parental leave rights by three weeks on April 1, 1992. As mentioned in Section 2, the 1993 reform was not a clean paternal quota reform: For parents of children born after April 1, the compensated parental leave period was extended by a total of seven weeks, *and*, from this date on, four (of the now 42) weeks of parental leave were reserved for fathers. Since our interest in the 1993 reform lies with the impact of only the paternal quota, we use the 1992 reform to remove the

[9] The children in our post-reform cohort will of course on average be somewhat younger than those in our pre-reform cohort. This may matter for the child outcomes we consider, as several studies have documented an association between season of birth and school performance (e.g. Strøm (2004) provides evidence for Norway). It is widely believed that this relationship is caused by differences in age at school entry, but it may also simply reflect that children born at different times of the year are conceived by women with different socioeconomic characteristics (Buckles and Hungerman (2008)). This age difference may also matter for some of the parental outcomes, as they are generally measured annually. Because mothers of children born before the reform, all else equal, will have a higher probability of a full year's income even several years later, one might spuriously attribute to the reform what is in reality a mere child age effect.

effect of the increase in general parental leave typically taken by the mother.[10]

We thus have four groups of parents: The 1992 pre- and post-reform cohorts, and the 1993 pre- and post-reform cohorts. Their respective parental leave rights are given in Table 1.

Table 1: Parental leave scheme in Norway 1992-1993

	Before April 1	After April 1	Difference
1992	32w, 0w pat. quota	35w, 0w pat. quota	3w gen. leave
1993	35w, 0w pat. quota	38w, 4w pat. quota	3w gen. leave + 4w pat. quota

3.2 Strategic timing of births

The 1993 reform was a large reform in the history of the Norwegian parental leave scheme. With the exception of the 1977 reform, all previous parental leave reforms meant an extension of four weeks or less. Seven weeks leave with full wage compensation is a considerable benefit at a time when childcare slots were rationed and many parents went on unpaid leave to care for small children. The 1993 reform therefore provided parents with strong incentives to have children born after April 1 rather than just before.

We see little reason to suspect that parents could time conception in anticipation of the reform. The national budget where the paternal quota was introduced became publicly available at October 7, 1992. At this time mothers who gave birth close to April 1, 1993 were already pregnant. Admittedly the reform itself was probably not very surprising to followers of the policy debate in Norway at the time, but there is little reason to expect that future parents knew the exact date of its implementation.[11] Searches in newspaper archives also suggest that the date of implementation was not publicly available before the national budget was presented.

Even if conception was not timed strategically, expecting parents with due dates close

[10] The 1992 reform was also announced during the autumn prior to its implementation; accordingly, there is little reason to fear that parents anticipated it and planned conception in order to fit the reform.

[11] Of the previous 7 parental leave reforms in Norway, implementation dates varied between April 1 (in 1989 and 1992), May 1 (in 1987 and 1990) and July 1 (in 1977, 1988 and 1991).

Table 2: Birth rate effects

	(1) ± 1 week	(2) ± 2 weeks	(3) ± 3 weeks	(4) ± 4 weeks
Panel A: Dependent variable is daily number of births				
Reform	18.0**	19.5***	9.00**	6.41*
	(7.54)	(5.39)	(4.39)	(3.79)
_cons	192.7***	202.6***	179.4***	181.7***
	(11.8)	(9.71)	(6.67)	(5.94)
Number of births moved	*63*	*136.5*	*94.5*	*89.7*
Observations	406	812	1218	1624
R^2	0.863	0.760	0.723	0.696
Panel B: Dependent variable is ln(daily number of births)				
Reform	0.099**	0.11***	0.051**	0.035
	(0.043)	(0.031)	(0.025)	(0.022)
_cons	5.27***	5.32***	5.18***	5.19***
	(0.067)	(0.055)	(0.038)	(0.034)
Share of births moved	*5.1%*	*5.7%*	*2.6%*	*1.8%*
Observations	406	812	1218	1624
R^2	0.866	0.766	0.730	0.706

Note: Sample is daily births within the relevant window (always centered around April 1), for the years 1975-2005. "Reform" is a dummy taking the value 1 for days in April 1993.

to April 1 could possibly postpone induced births or planned cesarean sections. Although the scope for strategic birth timing is limited since the vast majority of births in Norway are spontaneous vaginal deliveries, we investigate this claim empirically.[12] Following Gans and Leigh (2009), we run a regression where we relate the daily number of births to the reform (a dummy taking the value 1 for days after April 1 in 1993). We control for day of year fixed effects and for day of week fixed effects interacted with year fixed effects. In addition we add dummies for 10 days during Easter.[13] Our sample is daily births during the relevant time window (surrounding April 1) for the period 1975-2005, excluding 1989 and 1992 when parental leave reforms were implemented on April 1.

The analysis shows that day of week effects are considerable. This indicates that

[12] In 1993 the fraction of children born by cesarean section was around 12.4 percent, and of these deliveries, 59.4 percent were emergency operations. On average 12 percent of vaginal deliveries in 1993 were induced, while 88 percent were spontaneous (Folkehelseinstituttet, `http://mfr-nesstar.uib.no/mfr/`).

[13] In Norway, the Thursday and Friday before and Monday after Easter day are public holidays.

there is scope for non-medical reasons to affect the time at which children are born. As is reported in Table 2, we do find statistically significant evidence of strategic timing of births. The reform seems to have increased the daily number of births by 19.5 on average for the first two weeks of April relative to the last two weeks of March. This estimate implies that a total number of 137 births, or about 5.7% of the births predicted to have occurred in the last two weeks of March, were moved from somewhere in the latter half of March to somewhere in the first half of April 1993.[14] That the 1993 parental leave reform seems to have induced some parents to strategically time births is also documented by Brenn and Ytterstad (1997).

If strategic timing of births is related to (unobservable) characteristics that matter for the outcomes that we consider, this will bias our estimates of paternity leave. We address this potential problem by excluding births occurring during the two last weeks of March and the two first weeks of April.

3.3 Empirical specification

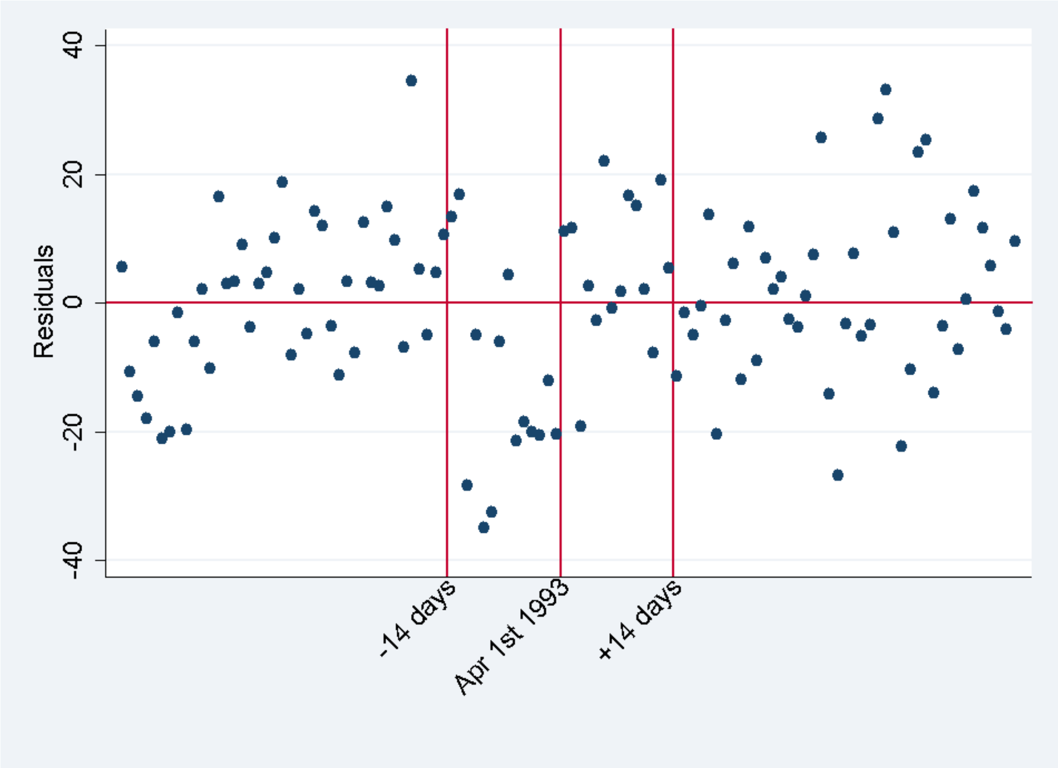
We estimate the following relationship based on data from families with children born in 1992 and 1993:

$$Y_i = \alpha PaternityLeave_i + \beta X_i + \delta_W Week_i + \delta_Y 1993_i + \epsilon_i, \tag{1}$$

where i is the child/household/parent indicator. Y denotes the parent, family or child outcome of interest to be discussed in Section 4, and X is a vector of pre-birth controls. $Week$ is a vector of dummies indicating during which week of the year the child was born. By including this vector we eliminate inherent differences between families with children born at different times of the year. 1993 is a dummy indicating whether the child was born in 1993 or in 1992. ϵ is the error term. α is the parameter of interest. For α to be

[14] Following Gans and Leigh (2009), the total number of births moved is calculated by dividing daily number of births by two (as one birth moved means one birth less in March and one more in April) and then multiplying by the number of days in the window. Similarly, the share of births moved is calculated by dividing the coefficient by two before converting log points to percentage points.

Figure 3: Daily births residuals



Note: Daily births residuals for the eight weeks centered around April 1, 1993 from a regression on day of year fixed effects, day of week fixed effects interacted with year fixed effects, and dummies for 10 days during Easter. Sample is based on data from daily births during eight-week time windows around April 1 for the period 1975-2005 (excluding 1989 and 1992).

given a causal interpretation we instrument $PaternityLeave$ with whether the child was born before or after the introduction of the paternal quota. The identifying assumption is that - other than the four weeks of paternal quota - there are no differences between the 1993 pre- and post-reform cohorts that do not also appear between the 1992 pre- and post-reform cohorts.

Our first stage is given by:

$$PaternityLeave_i = \rho Reform_i + \eta X_i + \gamma_W Week_i + \gamma_Y 1993_i + \nu_i, \tag{2}$$

where $Reform$ is a dummy variable taking the value 1 if the child was born after April 1, 1993. ρ is the regression adjusted compliance rate and ν is the error term.

Following Imbens and Angrist (1994), α is the *local average treatment effect* (LATE). This is the average treatment effect for *compliers*: that is the families whose treatment status (paternity leave) is affected by the paternal quota reform. An empirical description of this group is given in Section 5.4.

3.4 Eligibility and sample criteria

As mentioned in Section 2, both parents' right to paid parental leave is contingent on the child's mother having worked at least 50% during six out of the last ten months before the child's birth. Hence, families where mothers did not work the required amount were not covered by the paternal quota reform.

Since we do not perfectly observe eligibility status, we rely on parents' income history to capture this. In order to be considered eligible, both parents need to have an income above twice the 'basic amount' of the Norwegian social security system the year before the child was born (see footnote 5). 57% of all families fulfill this criterion. As our income criterion is rather strict, we may be excluding families that were actually eligible. With the strict criterion we are fairly confident, however, that our sample consists of families that were truly eligible for the paternal quota if they had children born after April 1, 1993. This is similar for the 1992 reform, as the rules for eligibility did not change with either reform.[15]

3.5 Time window

We face a trade-off between low bias and high precision when choosing the time window on which to estimate (1). In a narrow time window there is less chance that our main estimates are contaminated by omitted variables. A broader window would provide more precision by increasing the number of observations.

[15] Since information on both the 1992 and 1993 reforms became public in October of the previous year, and we use income data from that calender year to capture eligibility status, parents may have some scope to select into eligibility. In the data this does not seem to be a problem, all of our results are basically unaltered if we lag the eligibility criteria one year.

To balance these concerns we use a ±13 week window as our baseline in all specifications. We also report results on a ±7 week window. In line with the discussion of parents' strategic timing of births, we exclude observations from the two weeks before and the two weeks after April 1. But we also report results for our baseline window where these weeks are included.

4 Data

4.1 Child outcomes

Given their young age, there is limited register information on these children. We do however have data on school performance from administrative registers. In Norway, primary and lower secondary school (in total 10 years of schooling) are mandatory. At the end of lower secondary school students are graded. These grades matter for admission to upper secondary schools. Most grades are set by the student's own teachers; however, every student is also required to take a written exam, which is anonymous and graded by teachers from another school. To get an unbiased measure of student ability, e.g. avoid problems with relative grading, we focus on these latter exam scores. The exam subject is chosen randomly from the core subjects Norwegian, English and mathematics. Grades take integer values from one to six. For ease of interpretation grades are standardized and measured in units of standard deviations.[16]

The school performance sample is not identical to the samples for parental and family outcomes. The main reason is that for some children we do not observe an exam score.[17] The two samples are similar in terms of observables, both in terms of distribution and in terms of change around the introduction of the paternal quota.

[16] Exam grades have a standard deviation close to one, such that this standardization has limited impact, and the coefficients are close to the estimated effects also in units of grade points.

[17] Almost all children who continue to live in Norway will end up in the lower secondary school data. However, a small minority leave school early or late, these are excluded from our analysis, and about 4 percent of the students have no written exam score. In addition, families having multiple births will be represented with more than one observation in the analysis of children's outcomes.

Table 3: Summary statistics, labor market outcomes.

	Mean	SD
Fathers		
- earnings, 2-5	294371.1	(124846.6)
- earnings, 6-9	343035.1	(164772.1)
- full time, 2-5	0.78	(0.34)
- full time, 6-9	0.79	(0.35)
- part time, 2-5	0.80	(0.33)
- part time, 6-9	0.80	(0.34)
Mothers		
- earnings, 2-5	159553.8	(78809.2)
- earnings, 6-9	184908.3	(99916.3)
- full time, 2-5	0.41	(0.41)
- full time, 6-9	0.43	(0.42)
- part time, 2-5	0.57	(0.40)
- part time, 6-9	0.60	(0.40)
N	28344	

Note: Sample is children born during the 26 weeks surrounding April 1, 1993, excluding two weeks before and after April 1, divided into those born during the 13 weeks preceding the reform and those born during the first 13 weeks after the reform. 20 hours of work or more per week is classified as part-time, 30 hours or more is classified as full-time. Earnings are given in constant 1998 NOK.

4.2 Labor market outcomes

Statistics Norway provides data on yearly income going back to 1967 for the entire Norwegian population. Earnings are given in constant 1998 NOK, and are truncated above the 99th percentile. Data on employment status are obtained from Statistics Norway *Employment register* ("Arbeidstakerregisteret"), which contains data on all Norwegian employees.[18] This time series starts in 1993. Work hours are only reported in three broad categories: 1-19 hours, 20-29 hours and 30 or more hours. To measure labor supply we construct dummy variables capturing whether the individual work at least 20 hours (which we classify as part-time) or at least 30 hours (which we classify as full-time). If an individual is not registered with any employment or is defined as self-employed[19],

[18]This data set is used in several previous studies of the Norwegian labor market. Bratsberg and Raaum (2010), for example, use this data set to analyze how immigrant employment affect wages in the construction sector.

[19]If we observe that the individual has income from work that exceeds 1G, and in addition the individual's entrepreneurial income exceeds her income from employment, the individual is classified as self-employed.

his or her hours are set to zero.

To facilitate interpretation we rely on averages of labor market outcomes based on earnings and labor supply for multiple years. Such aggregation is also useful since it improves statistical power to detect effects that go in the same direction within a domain, without increasing the probability of a Type I error (Kling et al. (2007), Deming (2009), Almond and Currie (2010)). The averages are constructed by normalizing all outcome variables to have a zero mean and a standard deviation of one, and then averaging over these outcomes.

Table 3 shows the variation in the labor market data based on averages from when the child is 2-5 years old and 6-9 years old.[20] For families with children born in 1993 (1992) 'earnings 2-5' refer to average yearly earnings in the period 1995 through 1998 (1994 through 1997). We do not report results for earnings based on data from when the child is one year old, since most parents will be taking part of their leave during this year.

In our sample, 78% of fathers work full time, and an additional 2% work part time. These numbers are constant across child ages. Mothers work less: When the child is 2-5 years old, 41% work full time, while an additional 15% work part time. The numbers increase slightly at later child ages.

4.3 Family outcomes

Data on marriage, divorce and parity come from Statistics Norway's family and demography files. We investigate the impact of paternity leave on the following family outcomes: parents' total number of children 15 years after the reform (2008), their probability of divorce, the probability that the father has his next child with the same woman (conditional on having another child), child spacing and the number of days parent take leave if they have another child. Table 4 provide descriptive statistics.

For the family outcomes we also construct an index. Following Deming (2009), we

[20] This is a natural division as compulsory schooling in Norway starts at age six.

Table 4: Summary statistics, family outcomes.

	Mean	SD
Mother's parity	2.54	(0.85)
Father's parity	2.63	(0.96)
Prob(Divorced by child age 14)	0.21	(0.41)
Prob(Next child together)	0.87	(0.34)
Child spacing (years)	3.51	(1.80)
Father's leave next child (days)	24.8	(26.4)
N	17257	

Note: Sample is children born during the 26 weeks surrounding April 1, 1993, excluding two weeks before and after April 1, divided into those born during the 13 weeks preceding the reform and those born during the first 13 weeks after the reform.

decide whether an outcome is to be considered positive or negative and change the sign of negative outcomes. Hence, completed fertility (measured as parity in 2008) for mothers and fathers and the probability of the couple having their next child together enter positively, whereas the probability that the couple is registered with divorce 14 years later and the distance in time to their next child enter negatively.[21]

4.4 Control variables

We include control variables for parents' age at birth of their child and level of education and annual income the year before birth. We also control for birth order. Education is measured (October, 1) the year before birth, and is divided into four mutually exclusive categories; lower secondary education or less, upper secondary education, higher education lower degree and higher education higher degree. Birth order is controlled for by dummies for the number of children each parent already has, with six categories ranging from zero to five or more.

Table 5 give descriptive statistics for pre-birth characteristics for eligible parents whose children were born within a thirteen-week window prior or subsequent to April 1, 1993, excluding two weeks before and after April 1.[22] Columns (1) and (2) present averages

[21]Outcomes that, for obvious reasons, cannot be observed for a given family are excluded when generating the index.

[22]The characteristics are the same as those we control for in the regression. However, in Table 5 the

and standard deviations for the pre- and post-reform groups. Column (3) reports the estimated difference, and a test of equality for the two groups. Finally, column (4) presents a difference-in-difference estimate for each of the variables, comparing the differences between the pre- and post-reform groups with the corresponding differences for children born in 1992. Further descriptive statistics for the 1992 births are presented in Table 16 in the Appendix.[23]

The first thing to notice is how the 1993 reform changed the fathers' leave-taking behavior in our sample of eligible parents. Fathers' propensity to take parental leave increased by 36 percentage points, and their average number of leave days taken increased by 8 (meaning an average increase of 22.2 work days for those fathers actually taking leave - slightly more than the 4 weeks of paternal quota). This is in stark contrast to the year before. As the differences and diff-in-diff estimates are very similar, there was very little change in fathers' around April 1, 1992.

Also mothers' leave-taking behavior seem to have changed. Average leave days taken increased by 25, i.e. ten days more than the general increase in the parental leave (the increase not reserved for the paternal quota), and also about ten days more than the 1992 increase. This may be an indication that the paternal quota was enforced less strictly than the legislators intended, and that - at least initially - a fairly large share of the mothers got the paternal quota in addition to the rest of the leave. Furthermore, the *share* of mothers taking leave also increases markedly, by about six percentage points.

Having not found any change in regulations explaining the change in the share of mothers taking leave, this may be an artefact of changes in reporting practices that coincide with the introduction of the paternal quota, but we can not be sure about this. If the paternal quota was not taken by the fathers, but rather by the mothers, this could be expected to reinforce rather than change traditional roles in the household. However, while increasing paternal leave from zero to four weeks may produce a qualitative change,

categories for three, four and five or more older children have been combined for ease of exposition.

[23]The descriptive statistics presented here are based on the school performance sample. Descriptive statistics based on the sample of births are very similar.

Table 5: Descriptive statistics for cohorts born before and after April 1, 1993

	(1) Pre-reform		(2) Post-reform		(3) Difference		(4) Diff-in-diff	
	Mean	SD	Mean	SD	Estimate	SE	Estimate	SE
Fathers								
- % take leave	3.87	(19.3)	40.3	(49.1)	36.4***	(0.62)	35.9***	(0.68)
- no. leave days	2.02	(12.7)	10.3	(19.7)	8.29***	(0.27)	7.65***	(0.33)
- % age $<$ 25	4.15	(19.9)	4.66	(21.1)	0.51	(0.34)	0.53	(0.49)
- % age 25-29	27.6	(44.7)	28.5	(45.1)	0.90	(0.74)	1.08	(1.04)
- % age 30-34	36.0	(48.0)	37.4	(48.4)	1.39*	(0.79)	0.34	(1.11)
- % age $>$ 34	32.2	(46.7)	29.4	(45.6)	-2.80***	(0.75)	-1.94*	(1.06)
- % lower sec. or less	37.1	(48.3)	36.7	(48.2)	-0.46	(0.79)	0.34	(1.12)
- % upper secondary	34.8	(47.6)	34.4	(47.5)	-0.39	(0.78)	-1.56	(1.10)
- % higher ed. $\leq$ 4 yrs	19.3	(39.5)	20.2	(40.2)	0.96	(0.65)	1.30	(0.92)
- % higher ed. $>$ 4 yrs	8.79	(28.3)	8.68	(28.2)	-0.11	(0.46)	-0.079	(0.66)
- annual income	259.1	(100.2)	257.7	(99.7)	-1.36	(1.64)	-2.84	(2.27)
- % has no children	40.1	(49.0)	38.5	(48.7)	-1.55*	(0.80)	0.85	(1.13)
- % has one child	38.0	(48.5)	39.3	(48.8)	1.28	(0.80)	-0.50	(1.12)
- % has two children	16.1	(36.7)	16.8	(37.4)	0.74	(0.61)	0.40	(0.85)
- % has $\geq$ three children	5.84	(23.5)	5.37	(22.5)	-0.48	(0.38)	-0.75	(0.53)
Mothers								
- % take leave	90.4	(29.5)	96.2	(19.0)	5.87***	(0.40)	7.36***	(0.64)
- no. leave days	179.7	(67.9)	204.2	(68.5)	24.5***	(1.12)	10.2***	(1.54)
- % age $<$ 25	11.9	(32.3)	12.5	(33.1)	0.62	(0.54)	0.36	(0.77)
- % age 25-29	38.6	(48.7)	40.4	(49.1)	1.73**	(0.80)	1.99*	(1.13)
- % age 30-34	33.1	(47.1)	32.7	(46.9)	-0.41	(0.77)	-1.40	(1.08)
- % age $>$ 34	16.4	(37.0)	14.5	(35.2)	-1.94***	(0.59)	-0.95	(0.82)
- % lower sec. or less	38.3	(48.6)	36.1	(48.0)	-2.16***	(0.79)	-1.29	(1.12)
- % upper secondary	29.9	(45.8)	31.5	(46.4)	1.61**	(0.75)	1.24	(1.06)
- % higher ed. $\leq$ 4 yrs	27.8	(44.8)	28.2	(45.0)	0.43	(0.74)	0.67	(1.04)
- % higher ed. $>$ 4 yrs	4.03	(19.7)	4.14	(19.9)	0.11	(0.32)	-0.61	(0.45)
- annual income	178.2	(58.6)	176.7	(58.6)	-1.54	(0.96)	-2.97**	(1.34)
- % has no children	42.0	(49.4)	40.5	(49.1)	-1.50*	(0.81)	0.87	(1.14)
- % has one child	39.2	(48.8)	39.9	(49.0)	0.72	(0.80)	-0.93	(1.13)
- % has two children	14.8	(35.6)	16.0	(36.7)	1.18**	(0.59)	0.43	(0.83)
- % has $\geq$ three children	4.01	(19.6)	3.61	(18.7)	-0.40	(0.31)	-0.37	(0.43)
N	7203		7752		14955		30116	

Note: All observations except those regarding parental leave are taken from the year before the child's birth. Age categories are based on parents' age at birth of the first child. Sample is children born during the 26 weeks surrounding April 1, 1993, excluding two weeks before and after April 1, divided into those born during the 13 weeks preceding the reform and those born during the first 13 weeks after the reform.

the marginal effect of maternity leave, when it already is over 30 weeks, is likely to be much smaller.

Other than parental leave, there are few statistically significant differences in the pre-birth characteristics between the pre- and post-reform 1993 cohorts. Furthermore, these are largely matched in the 1992 data, such that there is only one variable which has a difference-in-difference significant at he 5% level, and two more at the 10% level. With 26 variables tested, this is about what we would expect if there were no systematic differences. Thus, on the whole Table 5 gives support to the idea of the reform providing exogenous variation along the relevant dimension (and only this one): Parental leave.

5 Results

We now present our estimates of the causal effects of paternity leave, instrumented by the Norwegian 1993 paternal quota reform. For brevity, we present only the LATE in question (α in Equation 1). Tables including the coefficients on covariates are available upon request.

Our first stage results (see Appendix Table 17) show that the excluded instrument ($Reform$) is a strong predictor of paternity leave: The regression adjusted compliance rate is 0.37 for our baseline specification.

Although the compliance rate is high and our instrument strong, 4 weeks of paternity leave may not be considered enough time to really impact long run outcomes. In so far as our point estimates point to zero effects, or no precise effects, this may not be taken as proof that paternity leave does not at all have an impact on these outcomes; it may also be that the reform is too small.

Yet, the perceived mechanism is not merely a direct link from 4 weeks spent at home by fathers to a change in children's school performance 15 years later, or to a penalization by employers for the time taken off from work. Rather, in line with Becker (1985, 1991) and the stated intentions of Norwegian policy makers, the relatively short period of paternity leave is assumed to affect the evolution of household roles and labor sharing, with a small change in initial comparative advantages yielding a larger impact on the degree of specialization in the longer run.

For every outcome, we run four different specifications. In Tables 6 through 11, column (1) shows the results from regressions on equation (1) on a ± 13 week window (excluding the ± 2 weeks that are affected by birth timing) without controls. In column (2) we have added the full set of family background variables available. This is our most preferred specification. Column (3) shows results when the time window is reduced to ± 7 weeks. Lastly, in column (4) we have included the ± 2 weeks affected by birth timing in the ± 13 week window.

When results are discussed without explicit reference to one particular specification, the specification in column (2) is the one in question.

5.1 Children's school performance

There is a rapidly growing literature on the importance of early childhood development for long term outcomes. Cunha and Heckman (2007) argue that skill formation early in life is determinant of skill development later on. Almond and Currie (2010), who have recently reviewed the literature in this field, state that child and family characteristics measured at the age of five "do as much to explain future outcomes as factors that labor economists have more traditionally focused on, such as years of education". Using Norwegian data, Carneiro et al. (2010) find that maternity leave significantly increases the probability that the child finishes high school. It is therefore of great interest to see how children are affected by paternity leave.

For the children born at the time of the 1993 paternal quota reform, there is naturally still a limited set of variables available. To measure cognitive skills we rely on pupils exam scores at the end of 10th grade (in 2009). 10th grade is the last year of mandatory schooling in Norway.

Table 6 presents the regression results for exam scores at the end of 10th grade.

Table 6: The impact of paternity leave on school performance at age 16

	(1) $\pm$ Weeks 3-13	(2) $\pm$ Weeks 3-13	(3) $\pm$ Weeks 3-7	(4) $\pm$ Weeks 1-13
Father takes leave	0.085	0.095*	-0.020	0.067
	(0.063)	(0.058)	(0.088)	(0.053)
Observations	28797	28797	13613	34256
R^2	0.005	0.173	0.172	0.173

Note: Each column provides results from an IV regression on equation 1. A year dummy and calender week fixed effects are included in all specifications. Included in specifications (2)-(4) are a set of socio-economic background characteristics explained in Table 3. Standard errors in parentheses. * $p < 0.10$,** $p < 0.05$, *** $p < 0.01$.

Paternity leave does not seem to have a robust effect on average school performance. For our preferred specification (column (2)), there is a positive effect of paternity leave on school performance close to 1/10 of a standard deviation, significant at the ten percent level. The result is similar (but not statistically significant) when the birth timing weeks are included (column (4)) and in the specification without controls (column (1)). When we reduce the time window the effect is relatively imprecise and close to zero (specification (3)).

Relative importance of parents

Using Swedish data, Liu and Skans (2010) find that the duration of general parental leave (typically taken by mothers) has no effect on average on children's school performance. They do however find positive effects for children of mothers with tertiary education, suggesting that parental leave strengthens the relationship between maternal education and school performance. This may also be relevant in our context.

However, the effects of maternal and paternal leave may be expected to differ. Maternal leave will tend to make unpaid leave paid, or replace non-parental care (e.g. kindergartens and grandparents) with maternal care. Maternal education may matter because of differences in change of behavior between high- and low-educated mothers and because of differences in the productivity of maternal time. This makes it relevant to study separate effects by maternal education. On the other hand, as mentioned in the introduction

Table 7: The impact of paternity leave on school performance at age 16: Interaction effects with parental educational groups

	(1) ± Weeks 3-13	(2) ± Weeks 3-13	(3) ± Weeks 3-7	(4) ± Weeks 1-13
Father takes leave				
- father highest educ ($F > M$)	0.33***	0.26**	0.17	0.21**
	(0.11)	(0.10)	(0.16)	(0.092)
- equal educ ($F = M$)	0.015	0.030	-0.13	0.0092
	(0.12)	(0.11)	(0.17)	(0.10)
- mother highest educ ($F < M$)	-0.080	-0.0014	-0.11	-0.019
	(0.098)	(0.089)	(0.14)	(0.083)
Observations	28797	28797	13613	34256
R^2	0.003	0.173	0.171	0.173
p-value $(F > M) = (F = M)$	0.055	0.13	0.19	0.15
p-value $(F > M) = (F < M)$	0.0054	0.053	0.18	0.063
p-value $(F = M) = (F < M)$	0.55	0.83	0.91	0.83

Note: Each column provides results from an IV regression on equation 1. A year dummy and calender week fixed effects are included in all specifications. Included in specifications (2)-(4) are a set of socio-economic background characteristics explained in Table 3. Standard errors in parentheses. * $p < 0.10$,** $p < 0.05$, *** $p < 0.01$.

to this section, *paternity leave* may impact the evolution of household roles and labor sharing in the home. If paternity leave sets off a dynamic where the father is more involved in his child and the mother becomes relatively less important, i.e. that motherly care to some extent is replaced by fatherly care, we should expect the effect of paternity leave to depend on parents' *relative* "skill levels" as parents. More specifically, we may expect to find a positive effect on cognitive skills when care from a highly educated father displaces that of a less educated mother.[24]

Table 7 reports heterogeneous effects of paternity leave according to whether the father has higher education than the mother, the parents have equally high education, or the mother has higher education than the father. The results are obtained by interacting one dummy for each possibility.[25] In our sample, 35.8% of students belong to the first

[24]If relative education is indeed what matters, we may also expect to find some sign of a more positive effect for highly-educated fathers, irrespective of the mother's education. However, this approach is likely to severely understate the potential effect of parental leave, because of the high correlation in parents' education.

[25]An alternative way of addressing this question is to estimate group-specific equations, which provides similar estimates. We prefer to interact the group indicators, as this increases precision and facilitates

group (F>M), 27.6% to the second (F=M), and 36.6% of students belong to the third group (F<M). Table 18 in the Appendix shows that the first stage results are fairly similar across the three groups - although the regression adjusted compliance rate seems to increase with the mother being relatively higher educated.

In the families where the fathers have the highest education level we find that paternity leave increases school performance with 0.26 of a standard deviation (statistically significant at the 5% level). The effect is fairly stable across samples, but it is not statistically significant at conventional levels with the ± 7 weeks time window. The estimated effect in families where mother's education is the longest is consistently negative, but not statistically significant. The last three rows of table 7 provide p-values from tests of equality of the estimated coefficients. For our most preferred specification, the hypothesis that the effect is the same for families where the father has higher education than the mother as for families where the mother has higher education than the father, is rejected at the 10% level. Our results are thus similar to the findings of Liu and Skans (2010), in that the effect of parental leave differ by the parents' education.

Daughters and sons

Tables 8 and 9 show the results from regressions run on separate samples according to the child's gender. We see that the effect of paternity leave on daughters' school performance is strong - 0.38 of a standard deviation - and statistically significant at the 5% level in families where the father is relatively higher educated. The corresponding effect on sons' school performance is much weaker - and not statistically significant.

There are several potential explanations for the differential gender effects. Different uptake is not one of them: The first stage results for the subsamples show that uptake does not differ between the groups. Fathers of sons are just as likely to take paternity leave as fathers of daughters.[26] Therefore, either paternity leave spurs a different dynamic

comparison of the group-specific coefficients.

[26]See Tables 19 and 20 in the Appendix.

Table 8: The impact of paternity leave on daughters' school performance at age 16: Interaction effects with parental educational groups

	(1) ± Weeks 3-13	(2) ± Weeks 3-13	(3) ± Weeks 3-7	(4) ± Weeks 1-13
Father takes leave				
- father highest educ ($F > M$)	0.47***	0.38**	0.26	0.33**
	(0.16)	(0.15)	(0.22)	(0.13)
- equal educ ($F = M$)	-0.033	-0.0012	-0.28	0.032
	(0.17)	(0.16)	(0.23)	(0.15)
- mother highest educ ($F < M$)	-0.048	0.012	0.040	-0.0080
	(0.14)	(0.13)	(0.19)	(0.12)
Observations	13989	13989	6670	16669
R^2	0.002	0.167	0.170	0.165
p-value $(F > M) = (F = M)$	0.034	0.080	0.094	0.13
p-value $(F > M) = (F < M)$	0.016	0.061	0.46	0.058
p-value $(F = M) = (F < M)$	0.95	0.95	0.29	0.83

Note: Each column provides results from an IV regression on equation 1. A year dummy and calender week fixed effects are included in all specifications. Included in specifications (2)-(4) are a set of socio-economic background characteristics explained in Table 3. Standard errors in parentheses. * $p < 0.10$,** $p < 0.05$, *** $p < 0.01$.

Table 9: The impact of paternity leave on sons' school performance at age 16: Interaction effects with parental educational groups

	(1) ± Weeks 3-13	(2) ± Weeks 3-13	(3) ± Weeks 3-7	(4) ± Weeks 1-13
Father takes leave				
- father highest educ ($F > M$)	0.20	0.14	0.070	0.11
	(0.15)	(0.14)	(0.22)	(0.13)
- equal educ ($F = M$)	0.065	0.017	-0.043	-0.046
	(0.17)	(0.16)	(0.25)	(0.15)
- mother highest educ ($F < M$)	-0.12	-0.025	-0.23	-0.038
	(0.14)	(0.12)	(0.19)	(0.11)
Observations	14808	14808	6943	17587
R^2	0.006	0.171	0.167	0.171
p-value $(F > M) = (F = M)$	0.56	0.56	0.73	0.44
p-value $(F > M) = (F < M)$	0.13	0.37	0.31	0.40
p-value $(F = M) = (F < M)$	0.42	0.83	0.56	0.97

Note: Each column provides results from an IV regression on equation 1. A year dummy and calender week fixed effects are included in all specifications. Included in specifications (2)-(4) are a set of socio-economic background characteristics explained in Table 3. Standard errors in parentheses. * $p < 0.10$,** $p < 0.05$, *** $p < 0.01$.

Table 10: The impact of paternity leave on fathers' labor market outcomes

	(1)	(2)	(3)	(4)
	$\pm$ Weeks 3-13	$\pm$ Weeks 3-13	$\pm$ Weeks 3-7	$\pm$ Weeks 1-13
Earnings, 2-5	-0.096	-0.065	-0.13**	-0.074*
	(0.060)	(0.042)	(0.063)	(0.039)
Earnings, 6-9	-0.032	-0.017	-0.034	-0.018
	(0.061)	(0.047)	(0.070)	(0.043)
Full time, 2-5	-0.011	-0.020	-0.054	-0.031
	(0.054)	(0.053)	(0.079)	(0.049)
Full time, 6-9	-0.033	-0.036	-0.052	-0.045
	(0.055)	(0.054)	(0.081)	(0.050)
Part time, 2-5	-0.017	-0.027	-0.037	-0.045
	(0.053)	(0.053)	(0.079)	(0.049)
Part time, 6-9	-0.035	-0.041	-0.066	-0.049
	(0.055)	(0.054)	(0.081)	(0.050)

Note: Each cell represents coefficients from an IV regression on equation 1. A year dummy and calender week fixed effects are included in all specifications. Included in specifications (2)-(4) are a set of socio-economic background characteristics explained in Table 3. Standard errors in parentheses. * $p < 0.10$,** $p < 0.05$, *** $p < 0.01$. The number of observations varies with outcomes and window, for the baseline window of $\pm$3-13 weeks where control variables are included (column 2) the number of observations is 27775, 27706, 27806, 27774, 27806 and 27774 for the respective outcomes.

in fathers' time use when taken with daughters as opposed to with sons, i.e. the parent's *behavioral response* is different according to the child's gender. Or, the effect of time spent with parents differs for boys and girls; there's a gendered difference in the *productivity* of parents' time.[27]

Only time use data would give a complete answer to which hypothesis is true - a behavioral effect or a productivity effect, or a combination of the two. The results of Morgan et al. (1988) suggest a behavioral effect, whereas Cascio (2009) and Havnes and Mogstad (2011) are suggestive of a productivity effect. Our data on labor market outcomes do however provide useful information on the time spent by parents doing market work.

Table 11: The impact of paternity leave on mothers' labor market outcomes

	(1)	(2)	(3)	(4)
	± Weeks 3-13	± Weeks 3-13	± Weeks 3-7	± Weeks 1-13
Earnings, 2-5	-0.25***	-0.15***	-0.16**	-0.13***
	(0.061)	(0.046)	(0.068)	(0.042)
Earnings, 6-9	-0.22***	-0.14***	-0.17**	-0.12***
	(0.061)	(0.050)	(0.075)	(0.046)
Full time, 2-5	-0.060	-0.014	0.0014	-0.0069
	(0.055)	(0.050)	(0.074)	(0.046)
Full time, 6-9	-0.077	-0.041	0.0046	-0.011
	(0.055)	(0.052)	(0.078)	(0.048)
Part time, 2-5	-0.12**	-0.087*	-0.14*	-0.079*
	(0.053)	(0.050)	(0.075)	(0.046)
Part time, 6-9	-0.15***	-0.12**	-0.15**	-0.091*
	(0.054)	(0.052)	(0.077)	(0.047)

Note: Each cell represents coefficients from an IV regression on equation 1. A year dummy and calender week fixed effects are included in all specifications. Included in specifications (2)-(4) are a set of socio-economic background characteristics explained in Table 3. Standard errors in parentheses. * $p < 0.10$,** $p < 0.05$, *** $p < 0.01$. The number of observations varies with outcomes and window, for the baseline window of ±3-13 weeks where control variables are included (column 2) the number of observations is 28307, 28271, 28320, 28289, 28320 and 28289 for the respective outcomes.

5.2 Labor market outcomes

Table 10 shows the results from regressions on equation (1), where the dependent variables are the normalized averages of earnings and labor supply for fathers over the years when the child is 2-5 years old and 6-9 years old, respectively.

Our estimated effects of paternity leave on fathers' labor market outcomes are negative for every outcome and across all time windows. This is consistent with the hypothesis that paternity leave increases fathers' involvement at home, but the estimates are (with a few exceptions) not statistically significant. The point estimates we document are however in line with the ones found by Rege and Solli (2010), indicating that the lack of any statistical significant effects in our analysis may be due to low statistical power. The identification strategy of Rege and Solli (2010) is based on the assumption that time trends in the earnings of fathers of children of different ages through the 1990s would be equal in the absence of the reform.

[27] We are thankful to Shelly Lundberg for pointing out this distinction to us.

Table 11 shows the corresponding results on mothers' labor market outcomes. Contrary to what would be expected in a simple Beckerian framework, mothers' earnings are negatively affected by paternity leave, both in the short run (child age 2-5) and in the long run (child age 6-9). In the baseline window, the effect is 0.15 of a standard deviation reduction in earnings, statistically significant at the one percent level. Point estimates and precision are fairly stable across time windows.

The estimated average reduction in earnings seems to be driven by a negative effect of paternity leave on mothers' working hours, more specifically the probability that mothers work part time or more. Though less precise, the estimated effects on work hours are comparable in size to the effects on earnings.

Our results on working hours, thus, provide no support for the hypothesis that paternity leave will cause households to specialize less in line with traditional gender roles: It seems, rather, that traditional household specialization is intensified. Furthermore, the conjecture that paternity leave causes general earning differentials between men and women to decrease is not substantiated empirically.

The adverse effects of paternity leave on women's labor market outcomes may be due to complementarities in mothers' and fathers' time. If fathers choose to spend more time at home and less in the market due to a family policy that strengthens the ties between fathers and children, so do mothers.

Daughters and sons

The results on children's school performance showed that paternity leave increased the importance of fathers' education for the school performance of daughters. In Table 12 we report the effects on parents' labor market outcomes in separate samples for families who had daughters and families who had sons. Again, the effects are much stronger in families who had daughters. For fathers' labor market outcomes, the point estimates are even positive (but far from statistically significant at conventional levels) in families who had boys. The results on mothers' labor market outcomes reported above, are driven by

Table 12: The impact of paternity leave on parents' labor market outcomes by child's gender

	(1) Fathers of girls	(2) Fathers of boys	(3) Mothers of girls	(4) Mothers of boys
Earnings, 2-5	-0.088	-0.033	-0.26***	-0.057
	(0.062)	(0.057)	(0.066)	(0.063)
Earnings, 6-9	-0.047	0.015	-0.25***	-0.032
	(0.069)	(0.064)	(0.072)	(0.069)
Full time, 2-5	-0.093	0.050	-0.046	0.017
	(0.077)	(0.073)	(0.072)	(0.069)
Full time, 6-9	-0.13	0.051	-0.086	-0.00077
	(0.078)	(0.075)	(0.076)	(0.072)
Part time, 2-5	-0.11	0.053	-0.13*	-0.046
	(0.077)	(0.072)	(0.072)	(0.069)
Part time, 6-9	-0.11	0.029	-0.21***	-0.048
	(0.078)	(0.075)	(0.075)	(0.071)

Note: Each cell represents coefficients from an IV regression on equation 1. A year dummy and calender week fixed effects are included in all specifications. Included in specifications (2)-(4) are a set of socio-economic background characteristics explained in Table 3. Standard errors in parentheses. * $p < 0.10$,** $p < 0.05$, *** $p < 0.01$.

the sample of families who had girls.

Our results are indicative of a gender difference in parents' behavioral response to paternity leave, which again may contribute to the effect on daughters' school performance in families where fathers are relatively higher educated than mothers. As gender is unrelated to unobserved pre-birth characteristics, it is unlikely that the estimated differences reflect differences in unobservables.

5.3 Family outcomes

Our analysis is completed by looking at how paternity leave affects a set of outcomes more directly connected to family life. In light of the somewhat surprising results found on labor market outcomes, family outcomes may be informative. For instance, if parents' time investment in family life are complementarities, this could show up in fertility and divorce rates.

Table 13 provides our results on family outcomes. In the baseline specification none of

Table 13: The impact of paternity leave on family outcomes

	(1) ± Weeks 3-13	(2) ± Weeks 3-13	(3) ± Weeks 3-7	(4) ± Weeks 1-13
Mother's parity	-0.028	-0.020	0.059	-0.054
	(0.054)	(0.044)	(0.066)	(0.040)
Father's parity	-0.032	-0.031	0.028	-0.052
	(0.060)	(0.046)	(0.069)	(0.042)
Divorce by child age 14	0.0082	0.0072	0.011	0.033
	(0.068)	(0.067)	(0.067)	(0.049)
Next child together	-0.024	-0.028	-0.025	-0.044*
	(0.028)	(0.027)	(0.039)	(0.025)
Child spacing	15.3	29.6	44.2	54.0
	(57.2)	(55.0)	(79.3)	(50.1)
Father's leave next child	0.37	0.48	1.66	1.04
	(2.35)	(2.32)	(3.27)	(2.09)
Index of outcomes	-0.058	-0.053	-0.0013	-0.080**
	(0.048)	(0.040)	(0.059)	(0.037)

Note: Each cell represents coefficients from an IV regression on equation 1. A year dummy and calender week fixed effects are included in all specifications. Included in specifications (2)-(4) are a set of socio-economic background characteristics explained in Table 3. Standard errors in parentheses. * $p < 0.10$,** $p < 0.05$, *** $p < 0.01$. The number of observations varies with outcomes and window, for the baseline window of ±3-13 weeks where control variables are included (column 2) the number of observations is 27790, 27790, 4339, 16652, 14541, 9231 and 27790 for the respective outcomes.

these are statistically significant. The lack of effects on family outcomes affects the scope for potential mechanisms through which paternity leave influences other outcomes. For example, the negative effects on mothers' earnings and employment could have followed from an increase in subsequent fertility. Our results indicate that this is not the case.

As referred in Lundberg and Rose (2004), several authors have reported that, in the United States, having a son relative to a daughter increases the likelihood that a marriage will remain intact. In light of this, it may be that our finding no average effects may conceal heterogeneous effects according to gender. Indeed, it turns out that the point estimates consistently have opposite signs in the two samples, although the differences are mostly statistically insignificant.[28]

5.4 Characterizing compliers

Since any instrumental variable strategy identifies effects only for the subpopulation affected by the instrument, external validity is always a concern (Moffitt (2005)). It is therefore useful to characterize families whose behavior was actually affected by the paternal quota reform - 'compliers' in the terminology of Imbens and Angrist (1994). If the compliant subpopulation is similar to the general population, the case for extrapolating estimated causal effects to the general population is stronger.

Table 14 presents the likelihood that a complier has a particular characteristic relative to the population of eligible families. This is obtained by taking the ratio of the first stage across a particular covariate group to the overall first-stage (Angrist and Pischke (2009), Angrist and Fernandez-Val (2010)).[29]

Fathers in the compliant subpopulation have somewhat higher education and age than the average for fathers in our sample of eligible families. When it comes to income and work experience, there are slightly fewer compliant fathers in both the lowest and highest

[28]The results can be seen in Table 21 in the Appendix.

[29]We report results for all covariates we use in our regression analysis. While 'annual income' and 'years of employment' enters linearly in our regression analysis, we have split them in quartiles for the purpose of characterizing the complier group.

Table 14: Complier characteristics ratios

	Father's	Mother's
Lower sec. or less	0.895	0.798
Upper secondary	0.996	1.012
Higher ed.. lower level	1.157	1.188
Higher ed.. higher level	1.063	1.243
Age 20-24	0.714	0.793
Age 25-29	1.015	1.018
Age 30-34	1.033	1.031
Age 35-44	0.984	1.064
Income quartile 1	0.805	0.688
Income quartile 2	1.13	0.888
Income quartile 3	1.115	1.143
Income quartile 4	0.963	1.245
Experience quartile 1	0.949	0.949
Experience quartile 2	1.102	0.964
Experience quartile 3	1.003	1.053
Experience quartile 4	0.951	1.04

Note: Each cell represents the ratio between the first stage estimate for the particular covariate group relative to the overall first stage. The ratios can be interpreted as the likelihood that the compliant subpopulation has a certain feature relative to the likelihood of that same feature among all eligible families. The sample is parents of children born in the 26 week period surrounding April 1, 1993 (minus the four weeks immediately around April 1), who were eligible for parental leave.

quartiles. Overall, fathers' characteristics in the compliant subpopulation are not very different from eligible families in general. For mothers' characteristics, the differences are larger: Compliant mothers have less than half the propensity to belong to the lowest income quartile in our sample of eligible families. They are also more highly educated. The difference regarding mothers' income between the compliant subpopulation and eligible families in general should be kept in mind when interpreting the effects reported above.

6 Conclusion

The proponents of paternity leave list a number of ways in which it will benefit fathers, children - and mothers. In this paper we look at the empirical basis for their claims.

We find no evidence of an effect of paternity leave on children's school performance on average. We do however find that the effect depends on the parents' relative education;

where children of fathers with higher education than the mother benefit from paternity leave. This is an indication that paternity leave causes a shift from motherly to fatherly care at home.

This shift is not mirrored in the parents' labor market outcomes, however. That mothers seem to respond to paternity leave by reducing labor supply is at odds with the standard economic framework emphasizing specialization in household vs. market work. Our analysis suggests that family-oriented policies, even if directed towards fathers, may be ill suited to reducing earnings differentials between men and women. Further investigation into the mechanisms behind the negative effects on mothers' earnings and labor supply is needed. One possible explanation is that mothers' and fathers' time at home are complements.

This paper also shows that family policies may work differently according to the child's gender, because parents respond differently to the policy depending on whether they have a son or a daughter. We have shown that the difference is not in terms of compliance with the policy, but rather that it lies in how the policy makes parents behave. This finding needs to be further investigated.

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7 Appendix

Table 15: Parental leave reforms in Norway

Date	Total parental leave	Compensation rate	Paternal quota
1.7.1977	18 weeks	100%	-
1.5.1987	20 weeks	100%	-
1.7.1988	22 weeks	100%	-
1.4.1989	24(30) weeks	100(80)%	-
1.5.1990	28(35) weeks	100(80)%	-
1.7.1991	32(40) weeks	100(80)%	-
1.4.1992	35(44.4) weeks	100(80)%	-
1.4.1993	42(52) weeks	100(80)%	4 weeks
1.7.2005	43(53) weeks	100(80)%	5 weeks
1.7.2006	44(54) weeks	100(80)%	6 weeks
1.7.2009	46(56) weeks	100(80)%	10 weeks
1.7.2011	47(57) weeks	100(80)%	12 weeks

Source: http://www.nav.no/rettskildene/Rundskriv/183541.cms.

Table 16: Descriptive statistics for cohorts born before and after April 1, 1992

	(1) Pre-reform		(2) Post-reform		(3) Difference	
	Mean	SD	Mean	SD	Estimate	SE
Fathers						
- % take leave	2.84	(16.6)	3.47	(18.3)	0.63**	(0.28)
- no. leave days	1.42	(9.99)	2.08	(13.5)	0.65***	(0.19)
- % age < 25	4.88	(21.6)	4.92	(21.6)	0.035	(0.35)
- % age 25-29	28.9	(45.3)	28.7	(45.2)	-0.23	(0.74)
- % age 30-34	36.1	(48.0)	37.2	(48.3)	1.03	(0.78)
- % age > 34	30.1	(45.9)	29.2	(45.5)	-0.84	(0.74)
- % lower sec. or less	37.8	(48.5)	36.9	(48.3)	-0.88	(0.79)
- % upper secondary	33.1	(47.1)	34.3	(47.5)	1.21	(0.77)
- % higher ed. $\leq$ 4 yrs	20	(40.0)	19.7	(39.8)	-0.28	(0.65)
- % higher ed. > 4 yrs	9.10	(28.8)	9.05	(28.7)	-0.046	(0.47)
- annual income	253.4	(97.4)	254.7	(95.9)	1.31	(1.57)
- % has no children	42.0	(49.4)	39.7	(48.9)	-2.33***	(0.80)
- % has one child	37.2	(48.3)	38.9	(48.8)	1.71**	(0.79)
- % has two children	15.6	(36.2)	15.9	(36.6)	0.36	(0.59)
- % has $\geq$ three children	5.20	(22.2)	5.47	(22.7)	0.26	(0.37)
Mothers						
- % take leave	90.2	(29.7)	88.7	(31.6)	-1.49***	(0.50)
- no. leave days	159.0	(60.2)	173.4	(70.5)	14.4***	(1.07)
- % age < 25	13.3	(34.0)	13.6	(34.3)	0.29	(0.55)
- % age 25-29	40.4	(49.1)	40.1	(49.0)	-0.29	(0.80)
- % age 30-34	31.8	(46.6)	32.8	(46.9)	0.99	(0.76)
- % age > 34	14.6	(35.3)	13.6	(34.2)	-0.99*	(0.56)
- % lower sec. or less	39.4	(48.9)	38.5	(48.7)	-0.87	(0.79)
- % upper secondary	29.0	(45.4)	29.4	(45.6)	0.44	(0.74)
- % higher ed. $\leq$ 4 yrs	28.0	(44.9)	27.7	(44.8)	-0.29	(0.73)
- % higher ed. > 4 yrs	3.56	(18.5)	4.28	(20.2)	0.72**	(0.32)
- annual income	174.2	(57.2)	175.7	(58.3)	1.41	(0.94)
- % has no children	44.2	(49.7)	42.0	(49.4)	-2.24***	(0.80)
- % has one child	38.0	(48.5)	39.5	(48.9)	1.54*	(0.79)
- % has two children	14.3	(35.0)	15.1	(35.8)	0.74	(0.58)
- % has $\geq$ three children	3.44	(18.2)	3.40	(18.1)	-0.038	(0.30)
N	7495		7666		15161	

Note: All observations except those regarding parental leave are taken from the year before the child's birth. Age categories are based on parents' age at birth of the first child. Sample is children born during the 26 weeks surrounding April 1, 1992, excluding two weeks before and after April 1, divided into those born during the 13 weeks preceding the reform and those born during the first 13 weeks after the reform.

Table 17: First stage results

	(1) ± Weeks 3-13	(2) ± Weeks 3-13	(3) ± Weeks 3-7	(4) ± Weeks 1-13
Father takes leave	0.37***	0.37***	0.36***	0.37***
	(0.0069)	(0.0069)	(0.010)	(0.0064)

Note: Each cell represents coefficients from an OLS regressions on equation 2. A year dummy and calender week fixed effects are included in all specifications. Included in specifications (2)-(4) are a set of socio-economic background characteristics explained in Table 3. Standard errors in parentheses. * p < 0.10,** p < 0.05, *** p < 0.01.

Table 18: First stage results by parental education.

	(1) ± Weeks 3-13	(2) ± Weeks 3-13	(3) ± Weeks 3-7	(4) ± Weeks 1-13
Father takes leave				
- father highest educ ($F > M$)	0.34***	0.34***	0.32***	0.34***
	(0.0065)	(0.0065)	(0.0092)	(0.0059)
- equal educ ($F = M$)	0.35***	0.35***	0.35***	0.35***
	(0.0069)	(0.0069)	(0.010)	(0.0063)
- mother highest educ ($F < M$)	0.38***	0.38***	0.36***	0.38***
	(0.0069)	(0.0069)	(0.010)	(0.0064)

Note: Each cell represents coefficients from an OLS regressions on equation 2. A year dummy and calender week fixed effects are included in all specifications. Included in specifications (2)-(4) are a set of socio-economic background characteristics explained in Table 3. Standard errors in parentheses. * p < 0.10,** p < 0.05, *** p < 0.01.

Table 19: First stage results by parental education - daughters.

	(1) ± Weeks 3-13	(2) ± Weeks 3-13	(3) ± Weeks 3-7	(4) ± Weeks 1-13
Father takes leave				
- father highest educ ($F > M$)	0.34***	0.34***	0.32***	0.34***
	(0.0093)	(0.0093)	(0.013)	(0.0084)
- equal educ ($F = M$)	0.35***	0.35***	0.35***	0.35***
	(0.010)	(0.010)	(0.015)	(0.0093)
- mother highest educ ($F < M$)	0.37***	0.37***	0.36***	0.37***
	(0.0100)	(0.0100)	(0.014)	(0.0091)

Note: Each cell represents coefficients from an OLS regressions on equation 2. A year dummy and calender week fixed effects are included in all specifications. Included in specifications (2)-(4) are a set of socio-economic background characteristics explained in Table 3. Standard errors in parentheses. * p < 0.10,** p < 0.05, *** p < 0.01.

Table 20: First stage results by parental education - sons.

	(1) ± Weeks 3-13	(2) ± Weeks 3-13	(3) ± Weeks 3-7	(4) ± Weeks 1-13
Father takes leave				
- father highest educ ($F > M$)	0.34***	0.34***	0.33***	0.35***
	(0.0091)	(0.0091)	(0.013)	(0.0083)
- equal educ ($F = M$)	0.36***	0.36***	0.35***	0.35***
	(0.0094)	(0.0094)	(0.014)	(0.0087)
- mother highest educ ($F < M$)	0.39***	0.39***	0.37***	0.38***
	(0.0097)	(0.0097)	(0.014)	(0.0089)

Note: Each cell represents coefficients from an OLS regressions on equation 2. A year dummy and calender week fixed effects are included in all specifications. Included in specifications (2)-(4) are a set of socio-economic background characteristics explained in Table 3. Standard errors in parentheses. * p < 0.10,** p < 0.05, *** p < 0.01.

Table 21: The impact of paternity leave on family outcomes by child's gender

	(1) Families who had girls	(2) Families who had boys
Mother's parity	0.032	-0.069
	(0.063)	(0.061)
Father's parity	0.023	-0.080
	(0.066)	(0.063)
Divorce by child age 14	0.13	-0.087
	(0.10)	(0.088)
Next child together	-0.081**	0.016
	(0.041)	(0.036)
Child spacing	91.9	-23.5
	(81.6)	(74.0)
Father's leave next child	5.78*	-3.59
	(3.25)	(3.26)
Index of outcomes	-0.071	-0.039
	(0.058)	(0.056)

Note: Each cell represents coefficients from an IV regression on equation 1. A year dummy and calender week fixed effects are included in all specifications. Included in specifications (2)-(4) are a set of socio-economic background characteristics explained in Table 3. Standard errors in parentheses. * p < 0.10,** p < 0.05, *** p < 0.01. The number of observations varies with outcomes and window, for the baseline window of ±3-13 weeks where control variables are included (column 2) the number of observations is 27790, 27790, 4339, 16652, 14541, 9231 and 27790 for the respective outcomes.