Paterity Leave*

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Abstract

This paper documents that a parental leave reform directed towards fathers causally impacts children's cognitive skills. School performance at age 16 improves, but only in families in which the father has higher education than the mother - implying that the effect of paternity leave depends on the care it displaces. Investigating data on parents' labor market outcomes, yields no evidence of a Beckerian shift in gender specialization in the household, as both parents tend to work less as consequence of the parental leave reform. All of our documented effects are driven by families with daughters. This indicates that parents' response to paternity leave depends on the child's gender.

Keywords: parental leave, child development, labor supply

JEL Classification: J13, J22, J24, I21

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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 the labor market. Politicians and policymakers in Northern Europe are strong believers that paternity leave strengthens women's position in the labor market, reduces the gender wage gap and promotes bonding between children and fathers.[1]

Wishing to alter traditional patterns of household specialization, politicians 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 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.[3]

In this paper we investigate how paternity leave impacts a broad range of outcomes using Norwegian register data. To handle the selection problem we use a parental leave reform, implemented on April 1, 1993, to evaluate the causal effects of paternity leave on children and parents. The main feature of the reform was the introduction of a four-week paternal quota. This reform caused a drastic change in fathers' leave-taking behavior.

Using a difference in difference approach, we find that *children's school performance* increases in families in which the father has higher education than the mother. This points to the crucial importance of the counterfactual: As paternal care becomes relatively more important, the effect will depend on both the quality of the paternal care and the maternal care displaced.

Consistent with our first finding, *fathers' earnings and working hours* are negatively affected by paternity leave. However, contrary to what one might expect, there are

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

[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.)

[3]Germany, for example, introduced in 2007 a two months paternity quota intended to provide incentives for parents to share home and market work equally ('Entwurf eines Gesetz zur Einführung des Elterngeldes' (Deutscher Bundestag Drucksache 16/1989, 20.06.2006).

strong and statistically significant negative effects on *women's labor market outcomes* of their spouse taking paternity leave - although this result must be interpreted with some caution.[4] Finally, paternity leave has no robust 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 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)). We therefore investigate whether the consequences of paternity leave 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, although the gender difference is not statistically significant. The difference persists for the labor market outcomes as well: Parents with reduced working hours and earning due to parental leave reform are the ones who have daughters.

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 only a handful studies aimed at identifying causal effects of paternity leave. These studies use Scandinavian paternal quota reforms and predominantly focus on parental outcomes.[5] We are not aware of any previous studies on how children's cognitive outcomes are affected by paternity leave.

The rest of the paper is organized as follows: Section 2 discusses the institutional setting. 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

[4] The reform used as the basis for our quasi-experiment also increased maternal leave taking in some families. This may be driving the negative effects on female labor supply.

[5] 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 of the estimates in both studies is very low. Rege and Solli (2010) do, however, find a negative impact of paternity leave on fathers' earnings in Norway. Kotsadam and Finseraas (2011b,a), using Norwegian survey data, find long-lasting effects of paternity leave on the division of household work within the family.

6 concludes.

2 Paternity leave in Norway

In Norway, wage compensated parental leave has been extended repeatedly since the 1970s, from the 18 weeks of leave with full wage compensation first granted in 1977 to 47 weeks in 2011.[6] Out of the total number of weeks, there has always been a share that the parents are free to share between them.

The parental leave scheme offers 100% wage compensation for both men and women (or they can choose 80% compensation and a longer leave period), but exceptions exist that are particularly relevant to men's right to full compensation. First of all, only parents who have worked 50% or more during at least six of the last ten months before the child's birth are eligible for wage compensated leave, and if the mother of a child does not fulfill this requirement, the right is also lost for the child's father.[7] If the mother fulfills the requirement but works part time (between 50 and 100%), the father's compensation rate is reduced accordingly.

Second, income compensation also reaches an upper bound of six times the *basic amount* (G) of the Norwegian social security system.[8] And third, until 2008, when self-employed individuals were granted rights to full compensation, the compensation rate for the self-employed was at 65% of their income.

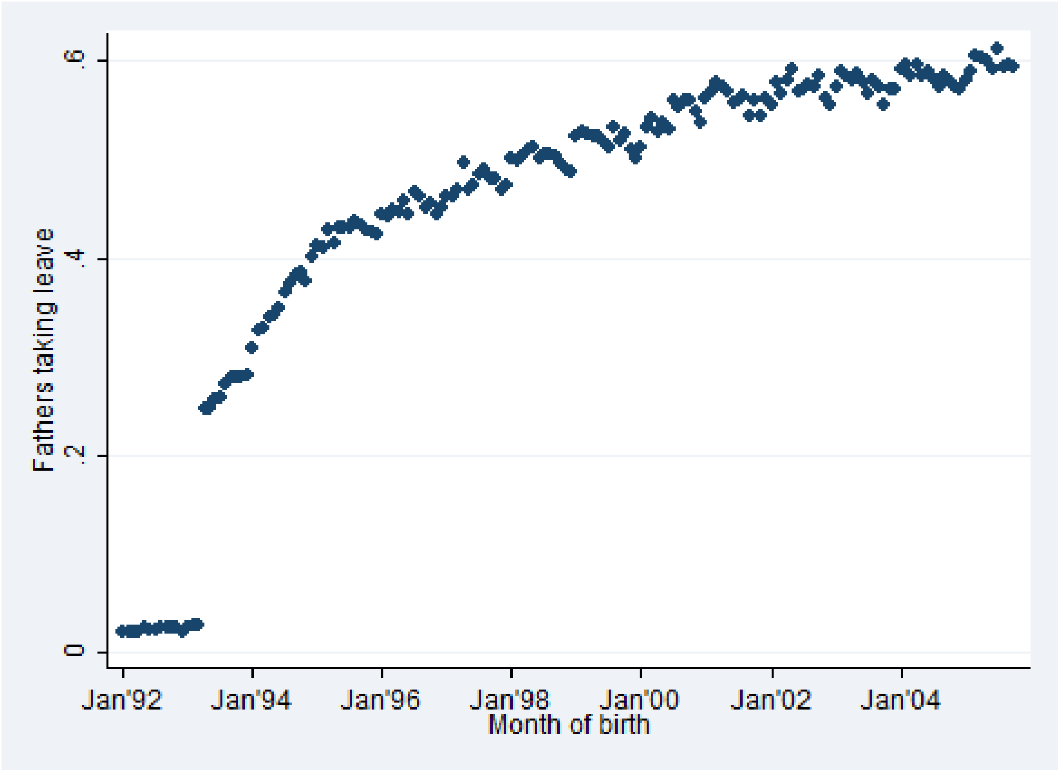
The fact that men tend to take very little of the leave period that can be freely shared between the parents triggered the labor party government to introduce a paternal quota in their suggestion for the national budget of 1993. The reform was passed in parliament in December 1992. Following implementation on April 1, 1993, four of the now 42 weeks of paid leave were to be reserved for the child's father. Barring "special circumstances",

[6]See Appendix Table 14 for a full description.

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

[8] *G* ("Folketrygdens grunnbeløp") 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).

Figure 1: Share of fathers taking leave, 1992-2006.



families would lose the right to these four weeks unless taken by the father. For the father to be eligible for the paternal quota the mother had to resume work.[9]

Figure 1, shows the fraction of fathers taking paternity leave, by the birth month of their child. There is a marked increase in the share of leave taking due to the reform, from 2.6% of fathers of children born in March 1993, to 24.6% of fathers of children born in April 1993. The share has continued to rise, and reached 60% in 2006. After the 1993 reform, every subsequent extension of parental leave in Norway has been fully absorbed in the paternal quota.

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

3 Identification: Reform as exogenous variation

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

This reform provides quasi-experimental variation in the uptake of paternity leave as long as there are no systematic differences between the families of children born before and after the reform date that also matter for the outcomes that we study. There are three potential pitfalls, to be discussed below.

3.1 Quasi-experiment not fully clean

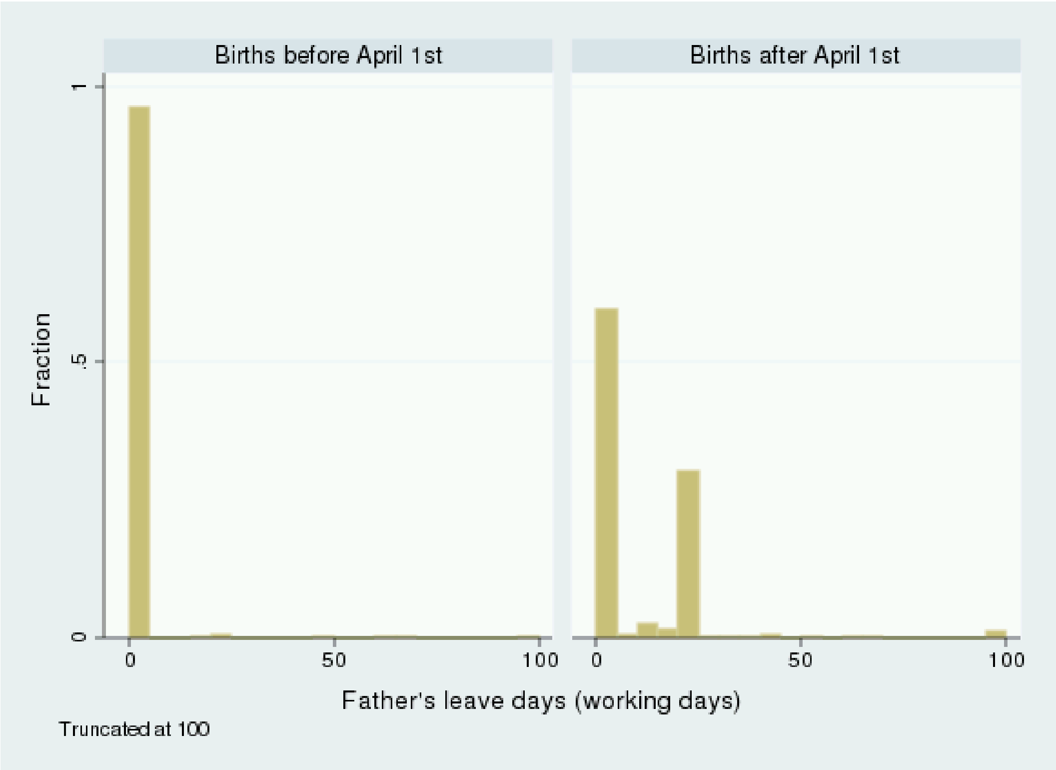
The 1993 reform combined two reforms: A four-week paternal quota reform and a three-week extension of the leave period that could be shared between parents.[10] In other words, there was a seven-week extension to a household's total parental leave, out of which four weeks were reserved for fathers (and three weeks were typically added to mothers' leave period).

Figure 2 gives the distribution of paternity leave spells for fathers of children born in a 26-week period surrounding April 1, 1993, in families who are eligible for parental leave.[11] The 40% percent of fathers eligible for the paternal quota that do take leave take on average approximately 25 days of leave (five weeks), with almost three quarters taking exactly the four weeks of the quota. Only 6% of all eligible fathers take more than

[10]There was also a further change, which we believe has less significance. Before the reform, mothers were obliged to start their leave period at the latest two weeks before their due date. After the reform this was increased to three weeks.

[11]As mentioned in Section 2, not all families were eligible for wage compensated parental leave. Therefore, all of our analyses are restricted to eligible families. How eligibility status is defined and determined will be discussed in Section 3.5.

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



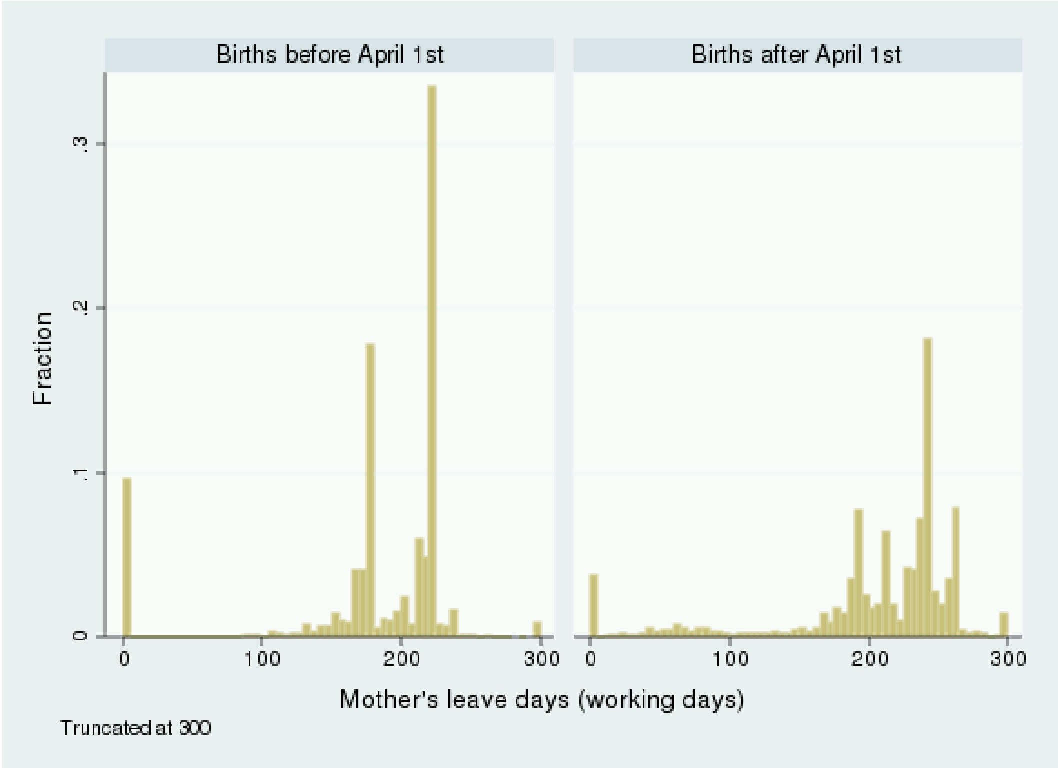
Note: The sample is fathers of children born in a 26-week period surrounding April 1, 1993, who were eligible for parental leave. Histogram bin width is five working days. For ease of exposure the number of leave days have been truncated at 100.

the quota, meaning that the 3 weeks of general parental leave extension were open to mothers.[12]

The fact that both parents' leave-taking behavior changes discontinuously at April 1 1993, strains our identification strategy. To address this problem we exploit a parental leave reform implemented exactly one year earlier. This parental leave reform included a three-week expansion of general parental leave. Based on the assumption that three weeks increase in general parental leave had similar effects in 1992 and 1993, a difference-in-differences approach, comparing the difference between the pre and post-reform cohorts in 1993 to that between corresponding cohorts in 1992, is suitable.

[12]Most fathers only have one spell of paternity leave - if any (the contingent probability is 90%). On average, their leave period starts when the child is nine months old. Less than 5% take leave after the child has turned one year.

Figure 3: Fraction of eligible mothers by number of leave days taken (working days).



Note: The sample is mothers with children born in a 26-week period surrounding April 1, 1993, who were eligible for parental leave. Histogram bin width is five working days. For ease of exposition the number of leave days have been truncated at 300.

An additional complication is that the paternity quota seems to have been enforced less strictly than the legislators intended. As mentioned above, families would normally lose the four weeks of parental leave unless they were taken by the father. However, exceptions could be granted by local caseworkers after a "special consideration of each individual case" (Ot. prp. nr. 13 (1992-93) p. 10).

Figure 3 documents that a fairly large share of mothers got the paternal quota in addition to the rest of their leave. Before the reform there are three spikes in maternal leave, one spike at zero, one spike at 35 weeks (reflecting maximum available leave at 100 % income compensation) and one spike at 44 weeks (reflecting 80 % income compensation). In addition there is some bunching at 33, 34 and 42, 43 weeks which reflects maximum leave taken, for a child born before the due date.[13]

[13]The total amount of leave days varies with the compensation rate the parents choose (80 or 100 %),

After the reform the spikes identified in the pre-reform data shift to the right by three and four weeks, for 100 % and 80 % compensation respectively. In addition there are two spikes at four weeks more than the expected maximum. The share of mothers taking up to four weeks more than the expected maximum days of leave is about 25 %. Thus, it seems that the paternal quota was transferred from the father to the mother in a substantial fraction of eligible families.[14]

The reform seems to result in one further change in the distribution of mothers' leave. More mothers seem to take at least some leave which is reflected in mothers taking very little leave (typically 50 to 100 days, i.e. much less than the days available). The fathers in these families have a somewhat higher propensity to take leave (about 50 %), and the mothers tend to have somewhat higher pre-birth income. This may indicate that the reform induced both parents in some relatively high-income families to take leave. Changes in reporting practices that coincide with the introduction of the paternal quota may also explain this finding. We have, however, not been able to find any documentation indicating that this is the case.

3.2 Differences related to birth season

The children in our post-reform cohort will 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)).

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the number of children born (multiple births give longer leave), and the extent to which the child is born before the due date (two (three) weeks of parental leave is potentially lost if the child is born early before (after) the reform).

[14]In the remaining 35 % of the families neither the mother nor the father took the paternal quota, indicating that it was foregone. This shows as a spike at minus four weeks in Figure 5 in the Appendix, which show the distribution of the sum of leave taken by the parents.

The age difference may also matter for the parents' 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, than mothers of children born after the reform, one might spuriously attribute to the reform what is in reality a mere child age effect. The difference-in-differences approach also addresses this problem. Using *week fixed effects*, we eliminate differences between families according to when their child is born.

3.3 Strategic timing of births

The 1993 reform was a large reform in the history of the Norwegian parental leave scheme. 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 in which the paternal quota was introduced became publicly available on 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.[15] Searches in newspaper archives also suggest that the date of implementation was not publicly available before the national budget was presented.[16]

Even if conception was not timed strategically, expectant parents with due dates close to April 1 could possibly postpone induced births or planned cesarean sections. A previous Norwegian study finds that mothers are able to influence their mode of delivery

[15] 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).

[16] The 1992 reform, implemented April 1 1992, 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.

Table 1: 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)
Constant	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)
Constant	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.

(Grytten et al. (2011)) so it is plausible that parents may also be able to influence the timing of a non-spontaneous birth.[17] Following Gans and Leigh (2009) we investigate this empirically.

We regress the daily number of births on a dummy variable for the reform (dummy equals one for dates after April 1, 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.[18] 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.

As is reported in Table 1, we find statistically significant evidence of strategic timing of

[17]The vast majority of births in Norway are spontaneous vaginal deliveries. In 1993 the fraction of children born by cesarean section was 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/`).

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

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. The 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.[19] 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.

3.4 Empirical specification

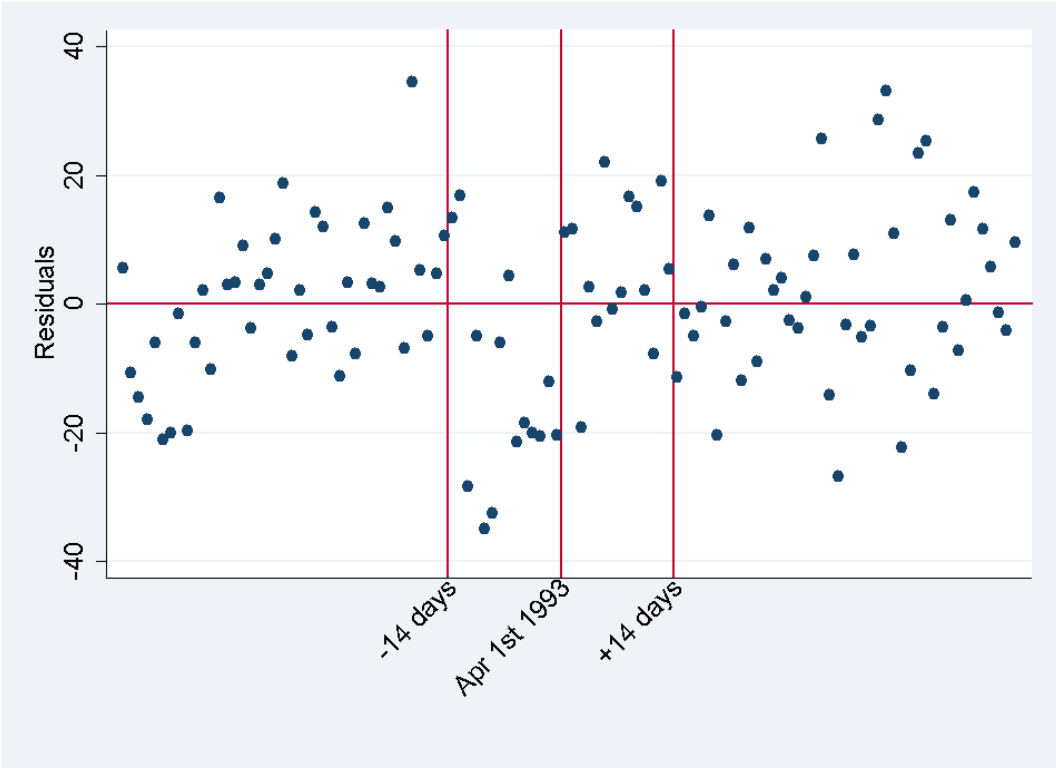
We estimate the following relationship based on data from families with children born in 1992 and 1993:

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

in which 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. This will be an intention to treat (ITT) estimate, i.e., the estimated effect of being exposed to the reform, irrespective of whether the father takes leave.

[19]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 4: Daily births residuals



Note: Daily births residuals for the eight weeks centered around April 1, 1993 from a specification including 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).

We have seen that the paternal quota increased both paternal and maternal leave taking. In families in which mothers took the paternal quota 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, the marginal effect of maternity leave, when it already is over 30 weeks, is likely to be much smaller. The ITT estimates are therefore arguably tapping the causal effects of paternity leave.

For child outcomes, the effect of paternity leave may depend on parents' *relative* "skill levels". In particular, this seems reasonable if paternity leave sets off a dynamic in which the father is more involved with his child and the mother becomes relatively less important, i.e. that maternal care to some extent is replaced by paternal care.[20] More

[20]This is in line with Becker (1985, 1991) and the stated intentions of Norwegian policy makers.

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.[21] For child outcomes we therefore report heterogeneous effects of paternity leave according to whether the father has higher education than the mother, the parents have equal levels of education, or the mother has higher education than the father.

3.5 Eligibility and sample criteria

As mentioned in Section 2, eligibility for paid parental leave is contingent on having worked 50% or more during at least six of the last ten months before the child's birth. Fathers' right to paid leave depends in addition on the child's mother having worked the required amount.

To avoid low precision in our estimates, it would be preferable to do estimations on a sample consisting of families who were actually affected by the reform. As we do not perfectly observe eligibility status, we use parents' income history to determine our sample. The trade off is between using a strict income requirement and excluding families from our sample who were, in fact, eligible, and using a less strict income requirement and including families who were not in fact eligible. The first type of error may affect the generalizability of our results, whereas the latter may give imprecise estimates.

We chose to use the middle ground requirement that for a family to be included in our sample both parents must have an income above twice the 'basic amount' of the Norwegian social security system during the year before the child's birth (see footnote 8).

57% of all families fulfill this criterion. Looking at the leave taking behavior of mothers (our most certain indication of eligibility), the criterion strikes a balance between excluding and including: In more than 90% of the families included by this criterion, the

[21]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.

mother is registered with paid maternity leave, whereas the number is only 33% in the families excluded. There is little to gain in terms of improving the former score by moving the income requirement up to three times the basic amount, as the share of mothers taking leave increases only marginally. There is, however, something to be lost, as 90% of the families thus excluded are registered with paid maternity leave.

Since information on both the 1992 and 1993 reforms became public in October of the previous year, and we use income data from that calendar year to capture eligibility status, there is some scope for parents to select into eligibility. All of our results are basically unaltered if we lag the eligibility criteria one year.

3.6 Time window

We face a trade-off between low bias and high precision when choosing the time window for estimating (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.

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 in which 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 (in 2009) from administrative registers.[22] In

[22] In June, 2012, the first children exposed to the 1993-reform will complete upper secondary schooling. In future work we intend to analyze results from upper secondary education.

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 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.[23]

The school performance sample is not identical to the samples for parental and family outcomes. This is because in the school sample, the unit of observation is each student completing compulsory schooling, while in the other samples, the unit is the family. We have some school results for about 95% of the total relevant cohorts of 16 years olds. For some of these children (about 5%) we do not observe an exam score. 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.

4.2 Labor market outcomes

Statistics Norway provides data on yearly "personal income" going back to 1967 for the entire Norwegian population. The personal income consists of wages, pensions and entrepreneurial income. In our analyses, 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.[24] 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

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

[24]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.

Table 2: Summary statistics, labor market outcomes.

	Mean	SD
Fathers		
- earnings, 2-5	294.4	(124.8)
- earnings, 6-9	343.0	(164.8)
- earnings, 10-14	390.2	(197.8)
- full time, 2-5	0.78	(0.34)
- full time, 6-9	0.79	(0.35)
- full time, 10-14	0.77	(0.36)
- part time, 2-5	0.80	(0.33)
- part time, 6-9	0.80	(0.34)
- part time, 10-14	0.78	(0.35)
Mothers		
- earnings, 2-5	159.6	(78.8)
- earnings, 6-9	184.9	(99.9)
- earnings, 10-14	237.0	(114.9)
- full time, 2-5	0.41	(0.41)
- full time, 6-9	0.43	(0.42)
- full time, 10-14	0.51	(0.43)
- part time, 2-5	0.57	(0.40)
- part time, 6-9	0.60	(0.40)
- part time, 10-14	0.67	(0.39)
N	28343	

Note: Sample is children born during the 26 weeks surrounding April 1, either in 1992 or 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 and are measured in 1000s.

measure labor supply we construct dummy variables capturing whether the individual is registered with at least 20 hours (which we classify as part-time) or at least 30 hours (which we classify as full-time) of employment per week. The dummy variables are set to zero otherwise. Due to imperfect observability, the analysis does not take working hours for the self-employed into account.[25]

To facilitate interpretation we rely on averages of labor market outcomes based on

[25] Working hours will therefore on average be somewhat underestimated. This is not a problem for our analysis as long as self-employment status does not depend on being treated by the reform. We have studied the impact of the reform on two different measures of self-employment, and we find a positive but generally not statistically significant effect on both mothers' and fathers' probability of being classified as self-employed. Our results on working hours are robust to excluding the self-employed thus defined from the analysis.

Table 3: Summary statistics, family outcomes.

	Mean	SD
Mother's parity	2.52	(0.84)
Father's parity	2.62	(0.92)
Prob(Divorced by child age 14)	0.20	(0.40)
Prob(Next child together)	0.87	(0.33)
Child spacing (years)	3.52	(1.82)
Father's leave next child (days)	24.9	(26.8)
N	28343	

Note: Sample is children born during the 26 weeks surrounding April 1, either in 1992 or 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.

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)). We construct the aggregated outcomes by normalizing all outcome variables to have a zero mean and a standard deviation of one, and then averaging over these outcomes.

Table 2 shows the variation in the labor market data based on averages from when the child is 2-5 years old, 6-9 years old and 10-14 years old. For families with children born in 1993 (1992) 'earnings 2-5' refers 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 stable across child ages. Mothers work less: When the child is 2-5 years old, 41% work full time, while an additional 16% work part time. The numbers increase slightly when the child is older.

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 breaking up (defined as either parent being registered as either divorced or a single parent at some point during the period from the child is 1 to 14 years old), 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 3 gives descriptive statistics.

4.4 Control variables

Table 4 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.[26] Columns (1) and (2) present averages 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.[27]

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 20.5 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

[26]The characteristics are the same as those we control for in the regression. However, in Table 4 the categories for three, four and five or more older children have been combined for ease of exposition.

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

Table 4: Descriptive statistics

	(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)
- % leave days $\geq$ 100	89.3	(30.9)	89.5	(30.6)	0.20	(0.50)	1.91***	(0.73)
- % 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. The sample used for column (1), (2) and (3) 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. In column (4) corresponding data for 1992 is used. For the difference and diff-in-diff estimates: * $p < 0.10$,** $p < 0.05$, *** $p < 0.01$.

change in fathers' leave-taking behavior around April 1, 1992.

Mothers' leave-taking behavior also 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 approximately ten days more than the 1992 increase. As discussed in Section 3.1, this probably reflects families in which the paternal quota is transferred to the mothers. The share of mothers taking leave also increases with the introduction of the reform. However, as there is little change in the share of mothers taking more than 100 days of leave, this seems to be driven by the fact that there are some mothers with very little leave after reform, as discussed in relation to Figure 3.

We include control variables for parents' age at birth of their child, their level of education and annual income the year before the child's birth, and the child's birth order. Education is measured on October 1 of the year before the child's 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.

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 generally what we would expect if there were no systematic differences. On the whole Table 4 gives support to the idea that the reform provides exogenous variation along the relevant dimension (and only this dimension): Parental leave.

5 Results

5.1 Characterizing compliers

Before moving to the main results it is useful to characterize families whose behavior was affected by the paternal quota reform - "compliers" in the terminology of Imbens and Angrist (1994). Table 5 presents the likelihood that a complier has a particular characteristic relative to the population of eligible families. This is obtained by running an auxiliary regression of the form (cf. Angrist and Pischke (2009), Angrist and Fernandez-Val (2010)):

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

First, we run this auxiliary regression for the entire population of eligible families. Unsurprisingly, given the descriptive statistics, we find that $Reform$ is a strong predictor of paternity leave: The regression adjusted compliance rate is 0.36 for our baseline time window. Then, for each of several subsamples, each defined by a variable in Table 5[28], we run the auxiliary regression and take the ratio to the overall compliance ratio. Table 5 provides descriptive statistics on both "male reform compliers" and "female reform compliers" (i.e. families in which the mother took the paternity quota).[29]

Both parents in the "male reform compliers" group have somewhat higher education, income and age than the average in our sample. Fathers' in this subpopulation are much less likely to be self-employed relative to eligible families in general. In terms of relative education within the family, there are only small differences, although in families in which the fathers have the highest relative education level the uptake rate is a bit lower than in the average eligible family.

The "female reform compliers" mirror the picture found for "male reform compliers".

[28] I.e. father has lower secondary or less education, mother has lower secondary or less, and so on.

[29] Female reform compliers are defined as mothers who, after April 1, 1993, are registered with maternity leave that is one to twenty days longer than the maximum length available (when not counting the paternity quota) and for whom the father of the child did not take paternity leave.

Table 5: Complier characteristics ratios

	Male compliers		Female compliers	
Regression adjusted compliance rate	0.361		0.224	
Parents' relative education:				
- father highest educ ($F > M$)	0.933		0.985	
- equal educ ($F = M$)	1.037		1.043	
- mother highest educ ($F < M$)	1.045		0.959	
Families who had boys	1.016		1.002	
	Fathers	Mothers	Fathers	Mothers
Lower sec. or less	0.915	0.799	1.129	1.198
Upper secondary	0.991	1.039	1.06	1.011
Higher ed., lower level	1.158	1.19	0.829	0.82
Higher ed., higher level	0.982	1.023	0.667	0.576
Age 20-24	0.748	0.813	1.437	1.398
Age 25-29	1.036	1.038	1.141	1.046
Age 30-34	1.025	1.004	0.945	0.862
Age 35-44	0.976	1.057	0.853	0.83
Income quartile 1	0.805	0.599	1.359	1.249
Income quartile 2	1.181	0.924	0.937	1.166
Income quartile 3	1.13	1.267	0.821	0.826
Income quartile 4	0.886	1.218	0.883	0.762
Self-employed	0.545	0.914	1.627	0.552

Note: The first row gives the overall regression adjusted compliance rate. The ratios reported in the other rows 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 in 1992 and in 1993 (minus the four weeks immediately around April 1), who were eligible for parental leave.

In these families both parents tend to be younger and less educated than the average eligible family. The father is much more likely to be self-employed than in the average eligible family. The regression adjusted transfer rate is 0.22.

5.2 Main Results

We now present our estimates of the causal effects of 1993 parental leave reform. For brevity, we present only the ITT in question (α in Equation 1). Tables including the coefficients on covariates are available upon request.

For every outcome, we run four different specifications. In Tables 6 through 10, column (1) shows the results from regressions estimating equation (1) for 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 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.

5.3 Children's school performance

In the upper panel of Table 6 regression results for exam scores at the end of 10th grade are presented. We find no significant effect on average school performance. For our preferred specification (column (2)), there is a statistically insignificant positive effect on school performance of about 3 percent of a standard deviation. The result is similar 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 imprecise and close to zero (specification (3)).

In the bottom panel of Table 6 we report heterogeneous effects of paternity leave according to whether the father has higher education than the mother, the parents have equal levels of education, or the mother has higher education than the father. In our sample, 35.8% of students belong to the first group (F>M), 27.6% to the second (F=M), and 36.6% of students belong to the third group (F<M). Across these three groups the reform affected paternity leave uptake similarly (as documented in the section 5.1).

In families in which the fathers have the highest education level we find that paternity leave increases school performance by 9% of a standard deviation (statistically significant at the 5% level). If this effect comes solely from the families in which the father actually does take leave, this amounts to an average effect of about 1/4 of a standard deviation

Table 6: Paterny leave and school performance at age 16

	(1)	(2)	(3)	(4)
	± Weeks 3-13	± Weeks 3-13	± Weeks 3-7	± Weeks 1-13
Average effect				
All students	0.031	0.034	-0.007	0.024
	(0.023)	(0.021)	(0.031)	(0.019)
Heterogeneous effects by parents' relative education				
- father highest educ ($F > M$)	0.114***	0.089**	0.057	0.073**
	(0.038)	(0.035)	(0.051)	(0.032)
- equal educ ($F = M$)	0.005	0.011	-0.045	0.003
	(0.044)	(0.040)	(0.058)	(0.037)
- mother highest educ ($F < M$)	-0.031	-0.000	-0.039	-0.007
	(0.037)	(0.034)	(0.050)	(0.032)
p-value $(F > M) = (F = M)$	0.063	0.140	0.190	0.151
p-value $(F > M) = (F < M)$	0.007	0.068	0.179	0.073
p-value $(F = M) = (F < M)$	0.534	0.835	0.944	0.829

Note: The top panel provides regression results for exam scores at the end of 10th grade. Each column provides ITT estimates from a regression based on equation 1. The bottom panel shows interaction effects with parental educational groups. A year dummy and calendar week fixed effects are included in all specifications. Included in specifications (2)-(4) are a set of socio-economic background characteristics explained in Table 4. Robust standard errors are in parentheses. The number of observations for the baseline window of ±3-13 weeks in which control variables are included (column 2) is 28797. * $p < 0.10$,** $p < 0.05$, *** $p < 0.01$.

in these families. The effect is fairly stable across samples, but it is not statistically significant at conventional levels with the smallest time window. The estimated effect in families in which mother's education is the longest is consistently negative, but not statistically significant. The last three rows of table 6 provide p-values from tests of equality of the estimated coefficients. For our preferred specification, the hypothesis that the effect is the same for families in which the father has higher education than the mother as for families in which the mother has higher education than the father, is rejected at the 10% level.

Heterogeneous results by maternal education are found also by Liu and Skans (2010), who study the duration of general parental leave (typically taken by mothers) and children's school performance. Using Swedish data, Liu and Skans (2010) find positive effects of parental leave, but only for children of mothers with tertiary education. We differ in studying the parents *relative* education. This reflects a difference in the likely counterfactual. When the total level of parental leave is extended as in Liu and Skans (2010), maternal care typically replaces formal or informal child care. In our case, paternal care is likely to replace maternal care. In this case it is parents relative education, rather than the absolute level of parental education, that matters.

Daughters and sons

Table 7: Paternity leave and school performance at age 16 - daughters

	(1)	(2)	(3)	(4)
	± Weeks 3-13	± Weeks 3-13	± Weeks 3-7	± Weeks 1-13
Average effect				
All students	0.045	0.046	0.007	0.042
	(0.032)	(0.029)	(0.042)	(0.027)
Heterogeneous effects by parents' relative education				
- father highest educ ($F > M$)	0.158***	0.127**	0.083	0.113**
	(0.054)	(0.050)	(0.071)	(0.045)
- equal educ ($F = M$)	-0.012	0.000	-0.097	0.012
	(0.063)	(0.057)	(0.083)	(0.052)
- mother highest educ ($F < M$)	-0.018	0.004	0.014	-0.003
	(0.051)	(0.048)	(0.069)	(0.044)
p-value $(F > M) = (F = M)$	0.041	0.092	0.102	0.143
p-value $(F > M) = (F < M)$	0.019	0.073	0.488	0.065
p-value $(F = M) = (F < M)$	0.936	0.956	0.306	0.826

Note: The top panel provides regression results for exam scores at the end of 10th grade. Each column provides ITT estimates from a regression based on equation 1. The bottom panel shows interaction effects with parental educational groups. A year dummy and calendar week fixed effects are included in all specifications. Included in specifications (2)-(4) are a set of socio-economic background characteristics explained in Table 4. Robust standard errors are in parentheses. The number of observations for the baseline window of ±3-13 weeks in which control variables are included (column 2) the number of observations is 13989. * $p < 0.10$,** $p < 0.05$, *** $p < 0.01$.

Tables 7 and 8 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 stronger - 13% of a standard deviation - and statistically significant at the 5% level in families in which the father has a relatively higher level of education. The effect in these families differs significantly from other families at the 10% level. If the estimated effect only comes from families in which the father takes leave, this corresponds to an average effect of approximately 38% of a standard deviation in these families. The corresponding effect on sons' school performance is much weaker - and

Table 8: Paternity leave and school performance at age 16 - sons

	(1)	(2)	(3)	(4)
	$\pm$ Weeks 3-13	$\pm$ Weeks 3-13	$\pm$ Weeks 3-7	$\pm$ Weeks 1-13
Average effect				
All students	0.015	0.015	-0.026	0.004
	(0.032)	(0.030)	(0.043)	(0.027)
Heterogeneous effects by parents' relative education				
- father highest educ $(F > M)$	0.069	0.048	0.024	0.036
	(0.054)	(0.049)	(0.073)	(0.045)
- equal educ $(F = M)$	0.023	0.006	-0.014	-0.015
	(0.062)	(0.056)	(0.082)	(0.052)
- mother highest educ $(F < M)$	-0.045	-0.010	-0.085	-0.015
	(0.053)	(0.049)	(0.072)	(0.045)
p-value $(F > M) = (F = M)$	0.579	0.577	0.729	0.451
p-value $(F > M) = (F < M)$	0.131	0.403	0.289	0.421
p-value $(F = M) = (F < M)$	0.406	0.825	0.516	0.994

Note: The top panel provides regression results for exam scores at the end of 10th grade. Each column provides ITT estimates from a regression based on equation 1. The bottom panel shows interaction effects with parental educational groups. A year dummy and calendar week fixed effects are included in all specifications. Included in specifications (2)-(4) are a set of socio-economic background characteristics explained in Table 4. Robust standard errors are in parentheses. The number of observations for the baseline window of ±3-13 weeks in which control variables are included (column 2) the number of observations is 14808. * $p < 0.10$,** $p < 0.05$, *** $p < 0.01$.

not statistically significant. However, the gender difference is not statistically significant. The estimated average effect, reported in the top panel of Tables 7 and 8, is somewhat larger for daughters than sons, but not statistically significant for either gender.

There are several potential explanations for differential gender effects. Different uptake is not one of them: The effect of the reform on uptake does not differ between the groups. Fathers of sons are just as likely to take paternity leave as fathers of daughters.[30] Therefore, either paternity leave spurs a different dynamic 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.[31] 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. Our data on labor market outcomes

[30]Results available upon request.

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

Table 9: Paternity leave and fathers' labor market outcomes

	(1)	(2)	(3)	(4)
	± Weeks 3-13	± Weeks 3-13	± Weeks 3-7	± Weeks 1-13
Earnings, 2-5	-0.035	-0.025	-0.047**	-0.028**
	(0.022)	(0.015)	(0.022)	(0.014)
Earnings, 6-9	-0.012	-0.008	-0.014	-0.007
	(0.022)	(0.017)	(0.025)	(0.016)
Earnings, 10-14	-0.012	-0.012	-0.004	-0.016
	(0.022)	(0.018)	(0.026)	(0.016)
Full time, 2-5	-0.004	-0.008	-0.020	-0.012
	(0.020)	(0.019)	(0.028)	(0.018)
Full time, 6-9	-0.014	-0.014	-0.019	-0.018
	(0.020)	(0.020)	(0.029)	(0.018)
Full time, 10-14	-0.006	-0.012	-0.004	-0.013
	(0.020)	(0.020)	(0.029)	(0.018)
Part time, 2-5	-0.006	-0.010	-0.014	-0.017
	(0.020)	(0.019)	(0.028)	(0.018)
Part time, 6-9	-0.014	-0.017	-0.024	-0.020
	(0.020)	(0.020)	(0.029)	(0.018)
Part time, 10-14	-0.006	-0.013	-0.007	-0.014
	(0.020)	(0.020)	(0.029)	(0.018)

Note: Each cell provides ITT estimates from a regression based on equation 1. A year dummy and calendar week fixed effects are included in all specifications. Included in specifications (2)-(4) are a set of socio-economic background characteristics explained in Table 4. Robust standard errors are 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 in which control variables are included (column 2) the number of observations is 27797, 27728 and 27645 for the respective earnings outcomes, and 27828 for the working hours outcomes.

do however provide useful information on the time spent by parents doing market work.

5.4 Labor market outcomes

Table 9 shows the results from regressions based on equation (1), in which the dependent variables are the normalized averages of earnings and labor supply for fathers for the years when the child is 2-5 years old, 6-9 years old, and 10-14 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

Table 10: Paternity leave and mothers' labor market outcomes

	(1) ± Weeks 3-13	(2) ± Weeks 3-13	(3) ± Weeks 3-7	(4) ± Weeks 1-13
Earnings, 2-5	-0.091***	-0.057***	-0.059**	-0.047***
	(0.022)	(0.017)	(0.024)	(0.015)
Earnings, 6-9	-0.080***	-0.051***	-0.063**	-0.046***
	(0.022)	(0.018)	(0.027)	(0.017)
Earnings, 10-14	-0.069***	-0.042**	-0.045*	-0.035**
	(0.022)	(0.018)	(0.027)	(0.017)
Full time, 2-5	-0.021	-0.005	0.001	-0.002
	(0.020)	(0.018)	(0.027)	(0.017)
Full time, 6-9	-0.029	-0.016	0.001	-0.005
	(0.020)	(0.019)	(0.028)	(0.018)
Full time, 10-14	-0.024	-0.015	-0.012	-0.010
	(0.020)	(0.020)	(0.029)	(0.018)
Part time, 2-5	-0.044**	-0.031*	-0.048*	-0.029*
	(0.019)	(0.018)	(0.027)	(0.017)
Part time, 6-9	-0.057***	-0.047**	-0.056**	-0.035**
	(0.020)	(0.019)	(0.028)	(0.017)
Part time, 10-14	-0.043**	-0.038**	-0.033	-0.025
	(0.020)	(0.019)	(0.028)	(0.018)

Note: Each cell provides ITT estimates from a regression based on equation 1. A year dummy and calendar week fixed effects are included in all specifications. Included in specifications (2)-(4) are a set of socio-economic background characteristics explained in Table 4. Robust standard errors are 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 in which control variables are included (column 2) the number of observations is 28315, 28280, and 28252 for the respective earnings outcomes, and 28328 for the working hours outcomes.

however consistent with those 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.

Table 10 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), the medium run (child age 6-9) and in the long run (child age 10-14). In the baseline window, the effect is 0.056 of a standard deviation reduction in earnings during the first years, 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.

Especially for the results on mothers' labor market outcomes, there is reason to worry that the causal mechanism goes not from paternity leave as identified with the paternal quota reform but rather through the increase in maternity leave that we document in Section 3.1. In order to assess this possibility, we investigate how results are affected by using 1991 as our comparison year instead of 1992. The main difference between these two years is that there was no parental leave reform during our time window in 1991. This means that using 1991 as a comparison year we do not difference out any effects of an increase in general/maternal leave. Hence, if maternity leave is what drives our results, they should become stronger when we use 1991. This is not the case.[32] It therefore seems unlikely that the changes in maternity leave are driving the results on mothers' labor market outcomes.

[32]Results available upon request.

Table 11: Paternity leave and 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	-0.036	-0.011	-0.094***	-0.022
	(0.022)	(0.021)	(0.024)	(0.023)
Earnings, 6-9	-0.020	0.006	-0.092***	-0.013
	(0.025)	(0.024)	(0.026)	(0.026)
Earnings, 10-14	-0.035	0.011	-0.075***	-0.011
	(0.026)	(0.024)	(0.026)	(0.026)
Full time, 2-5	-0.037	0.020	-0.016	0.007
	(0.028)	(0.027)	(0.027)	(0.026)
Full time, 6-9	-0.047*	0.018	-0.033	-0.001
	(0.028)	(0.028)	(0.028)	(0.027)
Full time, 10-14	-0.048*	0.024	-0.050*	0.017
	(0.029)	(0.028)	(0.028)	(0.027)
Part time, 2-5	-0.042	0.021	-0.046*	-0.017
	(0.028)	(0.027)	(0.026)	(0.026)
Part time, 6-9	-0.042	0.009	-0.075***	-0.022
	(0.028)	(0.028)	(0.027)	(0.026)
Part time, 10-14	-0.058**	0.031	-0.074***	-0.006
	(0.029)	(0.028)	(0.028)	(0.027)

Note: Each cell provides ITT estimates from a regression based on equation 1. A year dummy and calendar week fixed effects are included in all specifications. Included in specifications (2)-(4) are a set of socio-economic background characteristics explained in Table 4. Robust standard errors are in parentheses. * $p < 0.10$,** $p < 0.05$, *** $p < 0.01$.

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 11 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. The results on mothers' labor market outcomes reported above, are driven by the sample of families who had girls.

For fathers' labor market outcomes, we find statistically significant negative effects on labor supply in the long run (full time and part time work when the child is 10-14 years of age). The point estimates are even positive (but far from statistically significant at conventional levels) in families who had boys.

Our labor market results are indicative of a gender difference in parents' behavioral re-

Table 12: Paternity leave and family outcomes

	(1) ± Weeks 3-13	(2) ± Weeks 3-13	(3) ± Weeks 3-7	(4) ± Weeks 1-13
Mother's parity	-0.010	-0.007	0.021	-0.020
	(0.020)	(0.016)	(0.023)	(0.015)
Father's parity	-0.012	-0.010	0.013	-0.017
	(0.022)	(0.017)	(0.025)	(0.016)
Prob(Break up by child 14)	0.018*	0.018	0.013	0.017*
	(0.011)	(0.011)	(0.016)	(0.010)
Prob(Next child together)	-0.009	-0.011	-0.010	-0.016*
	(0.010)	(0.010)	(0.015)	(0.009)
Child spacing	5.862	11.690	18.319	20.824
	(21.942)	(21.280)	(30.509)	(19.435)
Father's leave next child	0.404	0.466	0.926	0.680
	(1.107)	(1.111)	(1.550)	(1.012)

Note: Each cell provides ITT estimates from a regression based on equation 1. A year dummy and calendar week fixed effects are included in all specifications. Included in specifications (2)-(4) are a set of socio-economic background characteristics explained in Table 4. Robust standard errors are 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 in which control variables are included (column 2) the number of observations is 27819, 27819, 27819, 16672, 14559 and 9238 for the respective outcomes.

sponse to paternity leave, which may have contributed to differential gender effects found for school performance. As gender is unrelated to unobserved pre-birth characteristics, it is unlikely that the estimated differences reflect differences in unobservables.

5.5 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 related to mothers' labor market outcomes, family outcomes could be informative. For instance, if parents' time investments in family life are complementary, this could show up in fertility and divorce rates.

Table 12 provides our results on family outcomes. In the baseline specification none of 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

Table 13: Paternity leave and family outcomes by child's gender

	(1) Families who had girls	(2) Families who had boys
Mother's parity	0.007	-0.022
	(0.023)	(0.023)
Father's parity	0.011	-0.031
	(0.024)	(0.023)
Prob(Break up by child aged 14)	0.023	0.013
	(0.016)	(0.015)
Prob(Next child together)	-0.030**	0.006
	(0.015)	(0.014)
Child spacing	32.976	-6.734
	(30.920)	(29.416)
Father's leave next child	2.895*	-1.577
	(1.502)	(1.619)

Note: Each cell provides ITT estimates from a regression based on equation 1. A year dummy and calendar week fixed effects are included in all specifications. Included in specifications (2)-(4) are a set of socio-economic background characteristics explained in Table 4. Robust standard errors are in parentheses. * $p < 0.10$,** $p < 0.05$, *** $p < 0.01$.

from an increase in subsequent fertility. Our results indicate that this is not the case.

Daughters and sons

As noted by Lundberg and Rose (2004), several authors have reported that, relative to having a daughter, having a son increases the likelihood that a marriage will remain intact in the United States. In light of this, it is possible that our finding of no average effects also might conceal heterogeneous effects according to gender. Indeed, we find that the point estimates generally have opposite signs in the two samples, although the differences are not statistically significant. In families who had girls, paternity leave decreases the probability of the parents having a next child together (contingent on the father having another child) by 3% (statistically significant at the 5% level). The results are reported in Table 13.

6 Conclusion

The proponents of paternity leave list a number of ways in which it will benefit fathers, mothers and children. In this paper we empirically examine these claims. Our empirical strategy exploits the fact that incentives to take paternity leave changed discontinuously on April 1, 1993. The strength of our approach is that time windows closely bracketing this date isolate plausible exogenous variation, allowing us to interpret our findings causally.

We find no support for the conjecture that family-oriented policies, even if directed towards fathers, are well suited for reducing earnings differentials between men and women, nor for the supposition that these policies contribute to increased fertility or lower divorce rates. Our analysis does, however, show that paternity leave causes fathers to be more involved in their children's lives. As a consequence of the parental leave reform, children's school performance improves in families in which fathers have higher education than the mother. This is an indication that paternity leave causes a shift from maternal to paternal care at home. The perceived mechanism is not merely a direct link from four weeks spent at home by fathers to a change in children's school performance 15 years later. Rather, consistent with Becker (1985, 1991) and the stated intentions of Norwegian policy makers, the relatively short period of paternity leave may affect the evolution of household roles, with a small change in initial comparative advantages yielding a larger impact in the longer run.

A limitation of our study is that only captures the *partial* effects of paternity leave on those affected by the paternal quota reform. There may also be important general equilibrium effects of fathers taking parental leave to an increasing extent. If, for instance, paternity leave leads employers to behave differently towards all men and women, this would not be captured by our empirical strategy. It would therefore be useful to complement our study with analysis utilizing alternative identification strategies.

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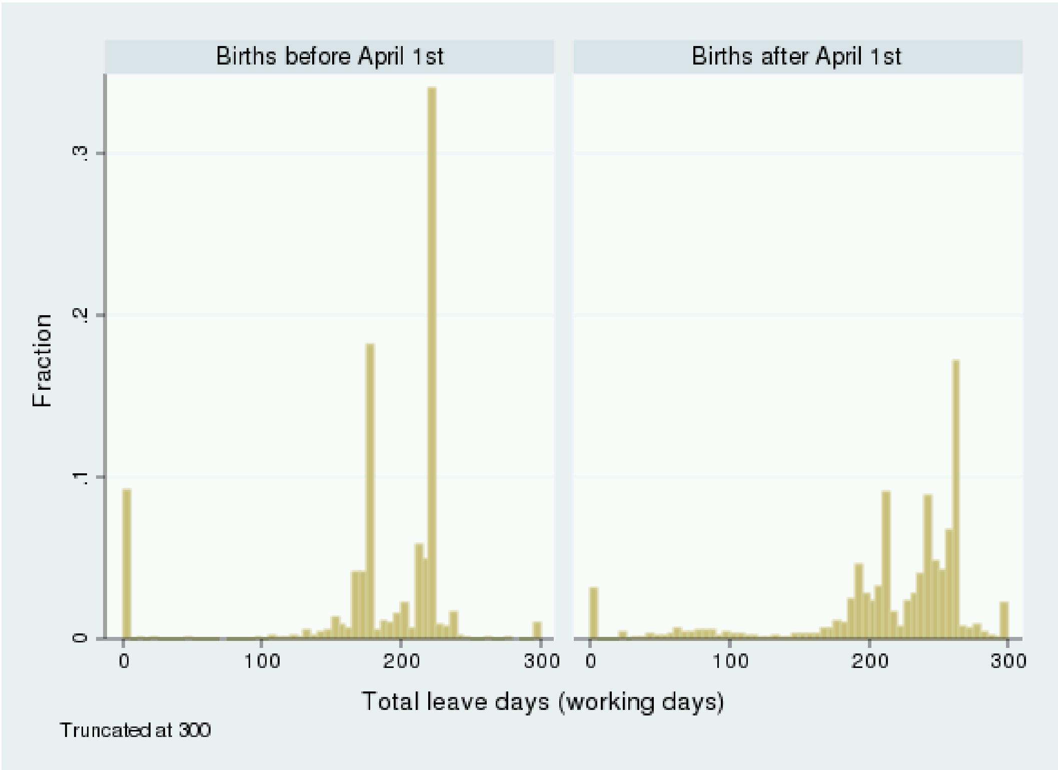
7 Appendix

Figure 5: Fraction of eligible families by number of leave days taken (working days).



Note: The sample is families with children born in a 26-week period surrounding April 1, 1993, who were eligible for parental leave. Leave taken is the sum of maternal and paternal leave. Histogram bin width is five working days. For ease of exposition the number of leave days have been truncated at 300.

Table 14: Parental leave reforms in Norway

Reform (date)	Parental leave (weeks)	Compensation rate	Maternal quota (weeks)	Paternal quota (weeks)
1.7.1977	18	100%	6	0
1.5.1987	20	100%	6	0
1.7.1988	22	100%	6	0
1.4.1989	24(30)	100%(80%)	6	0
1.5.1990	28(35)	100%(80%)	6	0
1.7.1991	32(40)	100%(80%)	8 (2 before birth)	0
1.4.1992	35(44.4)	100%(80%)	8 (2 before birth)	0
1.4.1993	42(52)	100%(80%)	9 (3 before birth)	4
1.7.2005	43(53)	100%(80%)	9 (3 before birth)	5
1.7.2006	44(54)	100%(80%)	9 (3 before birth)	6
1.7.2009	46(56)	100%(80%)	9 (3 before birth)	10
1.7.2011	47(57)	100%(80%)	9 (3 before birth)	12

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