Crowding out Dad?

The Effect of a Cash-for-Care Subsidy on Family Time Allocation*

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Abstract

This paper expands our understanding of possible specialization effects of extended parental leave policies. Identification is based on the introduction of the Cash-for-Care program in Norway in 1998, which increased mothers' incentives to withdraw from the labor market when their child was one and two years old. I estimate difference-in-differences models exploiting differences in individuals' exposures to the program among families with similar structures. Consistent with Schøne (2004) I find that the cash-for-care program decreased mothers' labor force participation by about four percentage points. Notably, however, I find no evidence that the fathers work more to compensate for the mothers declined labor supply.

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1 Introduction

In an attempt to assist parents in mitigating possible conflicts between work and family life, generous parental leave schemes are offered in many countries, with some countries even expanding the leave to which parents are entitled beyond one year into home care schemes.[1] Most analyses of these existing policies confirm their strong negative effects on the labor supply of mothers in the short run but suggest mixed effects in the longer run (Lalive and Zweimüller 2009, Schönberg and Ludsteck 2011, Drange and Rege 2012). At the same time, substantial evidence across several fields highlights the input of fathers as an important component of child development (Lamb et al. 1987, Harris et al. 1998, Tamis-LeMonda and Cabrera 2002). However, relatively little attention has been given to fathers when exploring the effects of new policies that encourage home care for a new child in a family.

One important implication of Becker's (1985) theory of specialization within the family is that if the mother allocates more time to home production, the father will increase his efforts in market production. Thus, if mothers indeed reduce their labor supply, as is the intention of a parental leave or home care scheme, fathers may have to work more to compensate for the potential income loss, particularly if the mother's income is not fully compensated by the scheme. Consequently, parental leave policies and, in particular, schemes that offer only partial income replacement could have the unintended effect of causing fathers to work more.

Importantly, if children indeed benefit from the father–child relationship, and fathers begin to work more in market production, and hence spend less time at home, these policies may in fact be detrimental for child development. This would particularly hold if the child had decreasing marginal returns from the inputs of both the mother and father. Thus, knowledge of how the specialization

[1] I refer to 'home care' schemes if parental leave extends beyond one year. Finland, Sweden, Norway, Austria and Germany (in some states) are among the countries offering these kinds of schemes (Schönberg and Ludsteck 2008, Lalive and Zweimüller 2009, Repo Rissanen and Sipilä 2010, Drange and Rege 2012).

pattern develops in families eligible for parental leave and home care schemes would improve our understanding of the consequences of such policies.

The aim of this paper is to investigate how the decline in labor force participation among mothers of young children affects labor force participation by fathers. The introduction of a Cash-for-Care subsidy in Norway allows me to study the effects of a universal home care subsidy (Cash-for-Care) on the allocation of time to market work by both parents in two-parent families.[2] The level of subsidy involved in this scheme was quite substantial, and thus may have replaced any lost income entirely in families in which the mother had low earnings potential. It is likely that the reaction of fathers in these families differed from those in families for which the subsidy did not fully replace the mothers' lost income. I will explore the potential heterogeneous effects of the subsidy by considering subsamples of families in which the mother had pre-reform earnings below or above median earnings.

Moreover, according to Becker's theory (Becker 1985), we expect to observe relatively less specialization in families in which the mother has a high investment in human capital. By comparing outcomes in a subsample of mothers that have completed a college degree with outcomes of mothers with lower educational accomplishment, I can empirically investigate this.

My analysis utilizes a comprehensive, longitudinal register database containing annual records for every person in Norway. I estimate difference-in-differences models by exploiting differences in individual exposures to the program across families with similar structures. Results are reported in terms of the subsidy's effect on employment and full-time attachment to the labor force, as well as for a more continuous measure: earnings. I consider similar outcomes for the analysis of fathers. Consistently with the findings in Schøne (2004), I show that the Cash-for-Care program decreased the attachment of mothers to the labor force by about four percentage points. Interestingly, mothers with low and high investment in human capital responded similarly to the policy change at the margin. The decrease in the full-time labor supply of mothers is consistent across groups with different levels of

[2] Several studies have shown that the subsidy caused a substantial decline in mothers' labor force attachment (Rønsen 2001, Schøne 2004, Drange and Rege 2012).

education. As for the impact on fathers, I find no evidence of any effect of the Cash-for-Care program on either workforce participation or earnings.

The remainder of the paper is structured as follows. Section 2 provides a brief overview of the literature on family specialization. Section 3 describes the Cash-for-Care reform program in Norway and gives a summary of other Norwegian policies directed toward families with young children. Section 4 presents the empirical strategy, Section 5 describes the data, Section 6 reports the empirical results, and Section 7 concludes.

2 Literature

Gary Becker developed an extensive theory of how family members specialize in market and home production (Becker 1965, 1985, 1991). In a household with two members, even a small initial difference in investment in a certain form of human capital (household or market human capital) will lead to specialization among household members. A household member who is relatively more efficient in market production will allocate more time to market activities, whereas a member relatively more efficient in home production will allocate more time to activities in the home. Typically, the arrival of children in the family implies more intensive home production, and thus at least one of the parents will allocate more time to the home.

We can consider the introduction of the Cash-for-Care subsidy in this analysis as a productivity shock to home production.[3] Leaning on Becker's theory of the division of labor in the household, the theoretical implication of the mother specializing in the home would be that the father allocates more time to market production. It is, however, also likely that the father will respond differently depending on the initial specialization of the mother. For example, if the mother did not previously work in paid employment, a cash subsidy would simply make household income higher. Given that leisure is a normal good, this may in fact cause the father to work less in market production. Alternatively, if the

[3] Studies using the same data source as this analysis (Rønsen 2001, Schøne 2004) suggest that the subsidy indeed caused a shift in the allocation of mothers' time away from market production.

mother worked full time before the introduction of the subsidy, and reduced her labor supply substantially as a response to the subsidy, she may have a lower income than before, even after receiving the subsidy. Hence, the family may be better off by allocating more of the father's time to work in market production in order to maintain the earlier level of income.

The existing empirical literature on specialization effects of parental leave or home care subsidies is sparse[4], but there are some related studies on household specialization when a child is born (regardless of subsidies). Angrist and Evans (1998) apply an instrumental variable approach to investigate the causal relationship between the arrival of a child in the family and the effect on the labor supply of the mother and father. In the analysis, they construct two instruments capturing exogenous variation in the number of children: one exploits the tendency of parents of two children of the same sex to have a third child more often than do parents of two children of different sexes, and another exploits the birth of twins.

Using these instruments, Angrist and Evans (1998) find a decline in maternal labor supply after a child is born. However, the decline in the labor supply is much smaller for women who have a college education and whose spouses have higher earnings. An additional child appears to have no significant effect on paternal labor supply. Consequently, Angrist and Evans (1998) conclude that if fathers spend more time at home when they have a third child,[5] it is only because they spend less time on leisure activities, not because they work less.

Moreover, while a number of studies have also investigated the effects of parental leave and home care schemes on the maternal labor supply, few have focused on the possible specialization effects.[6] An exception is Naz (2004). Utilizing data from the 1998 and 1999 living standard surveys to evaluate how the Cash-for-Care reform in Norway affected the specialization of couples in market and home

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[4] I am aware of studies investigating the effect of policies aimed at incentivizing the father to stay home with a new child. See for instance Rege and Solli (2010) or Cools et al. (2011).

[5] In their analysis, Angrist and Evans' instrument relies on the assumption that parents whose first two children are of the same sex are more likely to have a third child. Hence, variation in this variable affects the birth of the third child in the family, and so the causal interpretation of their findings is limited to parents with three children.

[6] Lalive and Zweimüller (2009) and Schönberg and Ludsteck (2011) only consider the effects of reform on the mother's labor supply.

production, Naz (2004) found that the subsidy decreased the labor supply of mothers but not of fathers. The present analysis differs from Naz (2004) in several respects. First, by employing register data, I can construct a data set with a substantially larger number of observations over several years. This means that I can observe household behavior over a longer period. In addition, I can evaluate whether households differ in their responses to the subsidy depending on their labor force attachment prior to program eligibility. Second, and equally importantly, as I have access to data after 1999, I can estimate the effect on parents of children fully included in the Cash-for-Care program. In contrast, Naz (2004) was obliged to consider children only partly affected by the program.[7] Nevertheless, despite these and several other key differences, my findings are largely consistent with Naz (2004).

3 Institutional setting

3.1 The Cash-for-Care program

In August 1998, the newly elected Christian Democratic government introduced the Cash-for-Care subsidy as one of the main pillars of its family policy. The government stated that the main goals of the subsidy were to ensure that families had more time to take care of their children, to allow families themselves to choose what kind of care they wished for their children and to equalize public transfers to families, regardless of the kind of care the family wanted for their child.

At the time of its introduction, the Cash-for-Care allowance constituted a significant part of family earnings, even for high-income families. The annual allowance was 36,000 Norwegian kroner (NOK),[8] and the average annual fee for publicly subsidized child care was about NOK 34,600 with some price subsidies for low-income families. Bettinger, Hægeland and Rege (2011) demonstrate that for a family in the bottom income quartile, the effective after-tax price of a full-time day care slot for a one- or two-year-old child constituted about 40 percent of average family earnings. For the third and

[7] Two- and three-year-olds in 1999 was partly treated as one- and two-year-olds in 1998. See Section 3 for details.
[8] The transfer was tax-free.

fourth income quartiles, the Cash-for-Care allowance constituted 15 and 10 percent of average family earnings, respectively.

Apart from the condition that a child subsidized under the Cash-for-Care program could not also attend publicly subsidized child care, the Cash-for-Care subsidy was unconditional. For instance, if neither parent wanted to stay at home with the child, they were free to hire a private child minder, or they could leave the child in the care of other family members. Thus, even if the parents of a child received the subsidy, this did not necessarily imply that one of the parents was taking care of the child during work hours. Moreover, parents were also free to receive parts of the subsidy, with the contingency being that the child would then have reduced hours in publicly subsidized child care.

Though the implementation of the Cash-for-Care program took place simultaneously throughout Norway, there was some variation in the starting times and ages of eligible children. From August 1998, all one-year-old children were eligible for the Cash-for-Care allowance, starting from the month after they turned one. From January 1999, both one- and two-year-old children were eligible.[9] Therefore, all children born from 1998 onward were eligible for 24 months of the Cash-for-Care allowance. For these children, eligibility started at the end or close to the end of maternity leave. We refer to these children as 'fully treated'. The Cash-for-Care allowance does not affect children born before 1996. However, children born in 1996 or 1997 could be eligible for as little as one month and as much as 24 months of the Cash-for-Care allowance. We refer to these children as partly treated.

Figure 1 describes the nature of the treatment. Each cell represents the age of a child in a given year. In the matrix, we can follow each cohort of children diagonally. The darkly shaded cells represent fully treated children; the lightly shaded cells, partly treated children. The value in each cell is the number of months for which the mother of a child of a given age in a given year was eligible for the subsidy. Note that some of the cells for older children are shaded even after the eligibility for the Cash-for-Care subsidy expired. This is because if the Cash-for-Care subsidy had a persistent effect on

[9] An exception to this rule was children who turned two after August 1, 1998. This ensured that no children had a break in the eligibility for the Cash-for-Care allowance.

the parental labor supply, we should discern a treatment effect in these cells. This is important to keep in mind when selecting the comparison group.

Figure 1: Months of eligibility

Age of child	1997	1998	1999	2000
One year	0	0–5	0–11	0–11
Two years	0	0–5	12	12
Three years	0	0	1–12	1–12
Four years	0	0	0	0
Five years	0	0	0	0

Notes: The darkly shaded area indicates the fully treated cohort; the lightly shaded area, the partly treated cohort. The first fully treated cohort comprised children born in 1998, i.e., who were one year old in 1999.

Figure 2 details the number of parents who received the subsidy (in full or in part) in the period 1998–2001. As shown, in 1999, the first year in which all families with one- and two-year-old children were eligible, three of every four families received a partial or full subsidy. It is also worth noting that the number of families receiving the subsidy differed only slightly between the two child age groups (one and two years): only about five thousand fewer families received the subsidy for a two-year-old child than for a one-year-old child.[10] It is also worth noting that very few fathers received the subsidy.

[10] We do not know if the same families received support both when their child was one and when the child turned two, but it is likely that most families that received the subsidy for a two-year-old child also received it when the child was one year old.

Figure 2: Families receiving the Cash-for-Care subsidy

	Recipients			Children			
	All	Fathers	Mothers	All	One-year-olds	Two-year-olds	% of all children aged 1–3 years
1998*	60,043	3,221	56,822	61,243	47,983	13,260	50.1
1999	86,224	3,743	82,481	89,592	46,598	42,994	74.8
2000	84,946	3,100	81,846	88,234	46,988	41,243	74.3
2001	84,169	3,008	81,161	87,580	46,549	41,031	73.2

Notes: * indicates the introduction of the subsidy for one-year-olds from August 1, 1998, and for two-year-olds from January 1, 1999. Source: Norwegian Welfare Administration.

3.2 Female labor market participation in Norway

In the decade prior to the introduction of the Cash-for-Care subsidy, there was a substantial increase in female labor market participation in Norway. For example, in 1991, 74 percent of married or cohabiting mothers with children aged 0–15 years were working, and this had increased to 81 percent when the Cash-for-Care subsidy was introduced in 1998 (Kjeldstad and Rønsen 2002). However, despite the high participation rate of Norwegian women, the incidence of part-time employment for women is above the OECD average, while the share of women with managerial responsibilities lies below the OECD average[11] (OECD 2011). Working women in Norway are also generally overrepresented in the public sector and, in particular, occupations relating to health and social work, where the prospect of obtaining part-time employment is better (Tronstad 2007).

The Norwegian government introduced the Cash-for-Care program at a time of extensive use of publicly subsidized childcare. At the time, about 40 percent of children aged one or two utilized publicly subsidized childcare,[12] and there was a shortage of places in these programs. At the time of the introduction of the Cash-for-Care subsidy, parents were entitled to 42 weeks of parental leave with

[11] OECD Family Database, based on statistics from 2007.
[12] OECD Labor Market Statistics (http://stats.oecd.org) and Statistics Norway (1998).

full wage compensation or, alternatively, 52 weeks with 80 percent wage compensation,[13] in addition to one year of unpaid job protection for each parent. On this basis, the Cash-for-Care program made it less costly for parents to extend the period they remained at home with the child before returning to work. However, if a mother chose to stay at home with her children until they turned three, receiving Cash-for-Care benefits, her job-protection period would have expired.

To receive the Cash-for-Care allowance, parents were required either to take care of the child themselves or to utilize informal care (e.g., relatives, neighbors or home-based day care). In Norway, formalized care consists almost exclusively of public and publicly subsidized private childcare centers. The same law regulates the two types of centers, and they basically offer the same type of program, have the same price schedule for parental pay and are equally subsidized. As very few private childcare centers did not run publicly subsidized programs, Cash-for-Care recipients in practice did not have the option of utilizing private formalized care.

3.3 Related family reforms

Before the introduction of the Cash-for-Care allowance in 1998, several work–family policies had already been implemented in Norway. In particular, there was a large extension in paid parental leave between 1986 and 1993. In 1986, Norwegian parents gained 18 weeks of paid parental leave, with leave rights subsequently extended to 35 weeks in 1992 and 42 weeks in 1993. Moreover, in 1993, Norway introduced a paternity quota for paid parental leave. This meant that of the 42 weeks of paid parental leave, four weeks were exclusively for the father. These policies began at least three years prior to the introduction of the Cash-for-Care allowance, and at least five years before the birth of the first cohort of children fully treated by the allowance.

For the most part, the uptake of the expansion in parental leave was immediate, while it was not until two years after implementation that extensive use of paternity leave began. This relatively slow

[13] In 2009, parental leave extended to 46 weeks of full compensation or 56 weeks of 80 percent compensation.

uptake of the paternity quota may raise concerns with the analysis, as it implies that the paternity quota fully affected the treatment group, whereas in the control group, the quota affected only the post-reform children. In terms of related empirical analysis, Rege and Solli (2010) show that the paternity quota affected the labor market attachment of fathers but not of mothers. Nevertheless, the following analysis pays careful attention to the introduction of the paternity quota.

Another relevant change at the time of the study was the 1997 school reform that reduced the school starting age from seven to six years and increased the period of mandatory schooling from nine to 10 years. Fortunately, all of the children in our sample started school at six years of age, so even if these children affected the labor supply of their mothers by starting school earlier,[14] the same effect should prevail for both the control and treatment groups. However, the school reform may still be of some concern because it led to an increase in the availability of publicly subsidized childcare slots in 1997, given that six-year-olds no longer needed childcare slots. This increase in childcare availability could possibly have increased the labor force participation of mothers while their children were young,[15] and could thereby bias our estimates downward. However, considering the development in childcare slots for the age groups in question, there appears to be little evidence of a spike in childcare attendance for five-year-olds in 1997, suggesting that childcare for this age group was not rationed at the time.[16] For the two-year-olds, however, there was an increase in enrolled children in 1997. This implies similar effects for both treatment groups (children born in 1995 and 1998) associated with the introduction of an early school start.

4 Empirical strategy

The introduction of the Cash-for-Care subsidy creates an exogenous variation in framework conditions facing parents of one- and two-year-old children before and after the policy reform. The

[14] Gelbach (2002) found that the maternal labor supply increases when the mother's oldest child starts school.
[15] See, for instance, Baker et al. (2008).
[16] Statistics Norway (2003).

difference-in-differences approach takes advantage of this by comparing changes in the labor supply of parents of two-year-olds pre- and post-reform with changes among parents of older children. Drange and Rege (2012) show that the effect of the subsidy on maternal labor supply was greatest when the child was two years old. Thus, if the subsidy indeed affects fathers, this is most likely to occur in the year in which the reduction in their spouses labor supply is largest. The main specification will hence investigate the effects for parents when they have a two-year-old child. The first fully treated cohort was born in 1998 (turning two in 2000), and the last cohort never treated was born in 1995 (turning two in 1997).

Let $ls_{2,00}$ be a dummy variable indicating whether a parent of a two-year-old in 2000 was employed. The difference-in-differences estimator can then be expressed by:

(1) $$\gamma_{2,00} = (ls_{2,00} - ls_{2,97}) - (ls_{5,00} - ls_{5,97}),$$

where $(ls_{2,00} - ls_{2,97})$ measures the change in labor force participation between 1997 and 2000 for parents of a two-year-old child. The purpose of the last term in Equation (1) is to control for trends in the labor market participation of parents with young children not affected by the reform (parents of five year olds). The difference-in-differences estimate will be positive (negative) if the Cash-for-Care reform has a positive (negative) effect on labor force participation, i.e., increasing (decreasing) participation.

The main analytical sample comprises parents whose youngest child was two or five years old in either 1997 or 2000. This ensures that parents of five-year-olds in the sample are not treated due to younger siblings. To make certain that the sample selection criteria are the same for the treatment and comparison groups, children are included only if they did not have a new sibling the year they turned five. This implies that while a mother who gives birth in both 1995 and 1998 is included with her two-year-old in 2000, she will not be included in the sample with the child born in 1995.

The main model specification is as follows:

$$\text{(2)} \qquad ls_{2,i} = \alpha + \beta age_{2,i} + \lambda year_{00,i} + \gamma_{2,00}(age_{2,i} year_{00,i}) + \eta X_i + \varepsilon_i,$$

where $ls_{2,i}$ is a dummy variable capturing the change of a given labor market outcome (for instance full time employed or not) of parent i with a two-year-old child (compared to five year-old child), $age_{2,i}$ is a dummy variable indicating that parent i has a two-year-old child (and not a five-year-old child), and $year_{00,i}$ is a year dummy variable equal to one if the outcome year is 2000 (and zero if the outcome year is 1997). The vector X_i captures a rich set of observable variables that may potentially influence the individual's labor supply choices (as discussed in Section 5). The parameter of interest in Equation (2) is γ. This captures, first, the effect on the labor supply of the interaction between having a child in the treatment group (a two-year-old relative to a five-year-old child), and second, whether the child actually belongs to the post-reform cohort (born in 1998 and fully treated by the reform).

Estimation of Equation (2) will produce unbiased estimates of γ only when the trends in labor supply for the two groups are similar. This identifying assumption may be difficult to defend for several reasons. In particular, we know that the mothers of young children in Norway substantially increased their level of labor market participation during the 1990s. If this trend differed between parents of two-year-olds and five-year-olds, the identifying assumption is violated. In Section 6 we investigate closely if trends seem to differ pre-reform.

Concurrent programs that affected the treatment and comparison groups differently may affect trends. Recall from the institutional setting presented in Section 3 that this period included the implementation of a number of other work–family policies. First, the introduction of the paternity quota influenced child cohorts born from 1993 onward, implying that this particular program did not affect the comparison group pre-reform (born in 1992). According to Rege and Solli (2010), the introduction of the paternity quota reduced paternal earnings, the effect being greatest when the child was two years old. Revisiting Equation 1, it becomes clear that if the fathers of children born in 1992 indeed worked more than did fathers with children born later (all other things being equal), the difference-in-differences approach will produce a positive treatment estimate of the Cash-for-Care subsidy, even if the subsidy did not affect fathers at all. Given that the paternity quota had the largest impact for two-year-old children, I will rely on including measures of the pre-Cash-for-Care paternal labor supply and earnings at this age for the comparison group as covariates in all analyses. If the effect estimates are robust to the inclusion of these covariates, this suggests that the estimates are not seriously biased by the introduction the paternity quota.

Among other policy changes that could confound findings, the introduction on February 1, 1995, of an additional year of job protection (one year for each parent, but without the possibility of transferring unused time to the other parent) is yet another candidate (NOU 1995). Like the paternity quota, this potentially affects all cohorts in the sample, except for the 1992 cohort. While it is unlikely that this policy change affected the labor supply of fathers directly, it could have affected mothers' work behavior and thereby exerted an indirect effect on fathers. Further, if the introduction of job protection induced mothers to work less post-reform, similar reasoning to that associated with the paternity quota would bias estimates for mothers upward. If fathers reacted to the reduced participation of mothers by working more, the estimate for fathers would be negative, given no effect of the Cash-for-Care subsidy. In a similar manner to the strategy implemented for handling the paternity quota, I will rely on including measures of the pre-Cash-for-Care maternal and paternal labor supply and earnings at this age for the comparison group as covariates in all analyses.

According to Hardoy and Schøne (2008), the Cash-for-Care subsidy also increased marital stability. This is a challenge to the empirical strategy because of the sample selection criterion that couples were still married when their youngest five-year-old child was included. For example, the composition of the sample could have changed if more people stayed married in the group of treated parents post-reform. It is unclear what effect this would have on the results. On the one hand, if increased marriage stability implies that more mothers with a better full-time earnings potential are included in the post-reform treatment group, there will be a downward bias in the estimates. On the other hand, if the mothers who remain married have a lower full-time earnings potential, the estimate will be biased upward. For this reason, I report robustness results constructed without the restriction on cohabitation status.

5 Data set description and restrictions

The analysis draws on the Norwegian registry data known as FD-Trygd as provided by Statistics Norway. This data set contains records on every Norwegian resident from 1992 to 2005, including individual demographic information (marital status, spouse identifier, sex, age, number of children), socioeconomic data (years of education, income, wealth), current employment status (full-time, part-time, minor part-time, self-employed), industry of employment, participation in welfare programs and geographic identifiers for county, municipality and neighborhood of residence.

The main analytical sample comprises couples with a child without a new sibling at five years of age. The child can have older siblings. Given that the idea is to explore possible specialization within the household, I restrict the sample to two-parent families. I define a two-parent family as the situation in which the mother and father of the child are cohabiting in the year the child turns five.

The data set also contains information on whether an individual works 4–19 hours, 20–29 hours or more than 30 hours per week. Based on this, I construct a variable that captures whether the mother or father is in any form of employment at the end of the year of evaluation, and a variable capturing whether the mother or father worked full time. However, because of lags in the submission of employee information by firms, some individuals are recorded as being in full-time employment despite the records also indicating very low or even zero earnings. I correct this by coding all individuals recorded as employed full time but who had incomes that seemingly precluded actual employment (i.e., very low or zero earnings) as unemployed.[17] It is worth noting that the information on work hours is only valid for people with an employer. Self-employed people are consequently also recorded as not being in full-time employment.[18] The variable capturing employment of any form therefore includes the three working-hour categories described above, and those not registered but with an income above ¼G for women (½G for men). I refer to this variable as *employed*. I define the dummy variable indicating full-time work as taking a value of one if the individual is working 30 hours or more per week and has earnings of more than 2G for women (4G for men), and zero otherwise. I refer to this variable as *full-time employed*.

The earnings variable captures all earnings that qualify for pensions, and is inflation adjusted with 1997 as the base year.[19] We code all missing observations on earnings as zero.[20] I do not specify the log of earnings as is sometimes typical in this sort of analysis, because I include all mothers in the earnings analysis, and some of these will not be working and will have zero earnings (log undefined). The reason I include all mothers in the earnings analysis is that the sample of working mothers is endogenous with respect to the Cash-for-Care reform.

The data allow for the construction of several variables capturing important child, father and mother characteristics to be included in the regression analysis. In order to ensure that the covariates are not endogenous to the reform, they are collected from a baseline year prior to the introduction of

[17] Every year, the Norwegian Labor and Welfare Administration defines a basic amount of money for the structuring of the Norwegian pension system, referred to as G. To correct the full-time employment variable, I define women earning less than 2G as out of full-time employment, while men earning less than 4G are defined as out of full-time employment (on average, women earn substantially less than men).

[18] I do, however, run the regression on a sample that includes self-employed people, and obtain very similar results. See the Appendix for details.

[19] I inflation adjust earnings for mothers to their 1997 level using the change in the earnings of the entire female population aged 20–67 years.

[20] The results are also robust after excluding these observations.

the Cash-for-Care subsidy. Accordingly, for children aged two in 1997 and 2000, I collect the covariates from the year prior to their birth year (1994 and 1997, respectively). For children aged five in 1997 and 2000, the covariates are from when they were two years old (1994 and 1997, respectively).

The control variables include the child's sex[21], the number of children (0, 1, 2, 3, 4, ≥5),[22] the mother's age (years) at the birth of the youngest child (<20, 20–24, 25–29, 30–34, 35–39, 40–44, ≥45), the mother's age (years) at the birth of the oldest child (<20, 20–24, 25–29, 30–34, 35–39, 40–45, ≥45) and the father's age (years). They also include education (completed high school, completed college), linear and quadratic controls for earnings, employment status (minor part-time, part-time or full-time) and indicators identifying the receipt of any social welfare benefits, living in a densely populated area (city), immigrant status, municipality-specific unemployment rates interacted with the birth year of the child and, finally, municipality fixed effects.

6 Empirical results

6.1 Summary statistics

Figure 3 depicts the trends in the share of full-time working mothers and fathers of children in the age groups included in the analysis.[23] As shown, we can discern a decline in the share of full-time working mothers starting in 1998, and this is suggestive of the effect of the Cash-for-Care reform on maternal labor force participation. However, we should note that only children born late in the year were eligible in 1998, and that 1999 is the first year that two-year-olds were eligible for assistance for the entire year. As expected, from 1999 onward, the full-time attachment gap between mothers of two-

[21] Lundberg and Rose (2002) find that fatherhood significantly increases labor supply and wage. Importantly, they find that labor supply increases more in response to the birth of a son than to the birth of a daughter.
[22] Parenthetical documentation on any control variable indicates the ranges of the series of categorical variables that characterize the specific trait.
[23] Because of a change in the registration procedure at Statistics Norway, there is a surge in missing values for the labor supply variable in 1998 and 1999. The share of missing values is similar across mothers of children of different ages and, according to Statistics Norway, is because of a change in registration routines during these years. See the Appendix for details.

and five-year-olds increases sharply. Prior to the reform, trends in the labor force participation of mothers are similar across the different age cohorts, up to the last year before the introduction of the subsidy.

Turning to the details for fathers, in Panel B of Figure 3 it appears that the trends in full-time attachment are similar, regardless of the age of the children. It is also likely that the decline in both age groups after 1998 relates to the abovementioned change in registration procedures. On this basis, there would appear to be little evidence of any change in the labor supply of fathers corresponding to the clear decline we observe for mothers.

Figure 3: Trends in parents' full time attachment

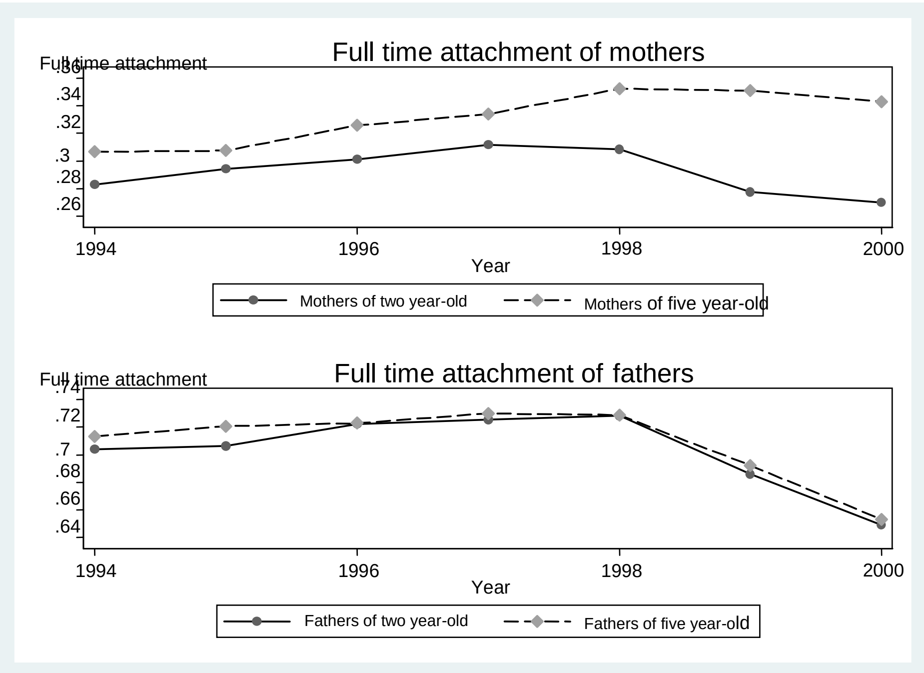

Figure 4 provides summary statistics on the earnings of mothers and fathers. As shown, from 1998 onward, the trends in the earnings of mothers of two- and five-year-olds begin to diverge, in line with what we have already observed concerning the likelihood of full-time work in Figure 3. The earnings trends for both mothers and fathers are also very similar before 1998, suggesting that the parents of five-year-olds are a suitable comparison group when implementing the difference-in-differences model. For fathers of five-year-olds, there appears to be a small decrease in earnings in 1998, and this is not reflected by a similar decrease for the fathers of younger children, resulting in a slightly smaller difference in earnings levels in that year. However, in 1999, earnings appear to revert to their pre-1998 levels, suggesting that it is unlikely that the Cash-for-Care subsidy can help explain the difference.

Figure 4: Trends in parents' earnings

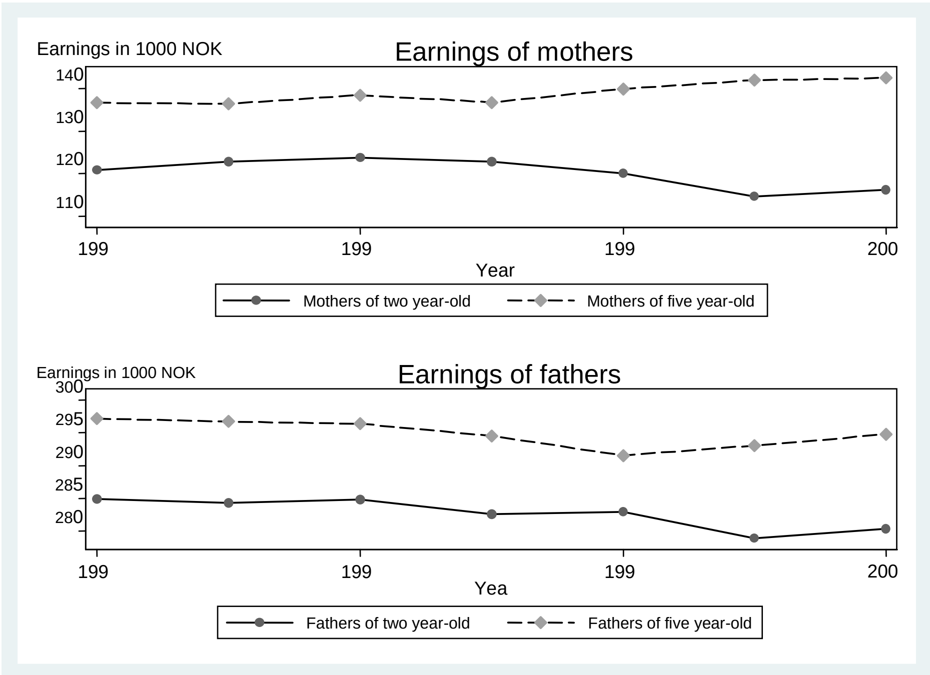

Table 1 details similar summary statistics. The outcomes for the mothers and fathers of treated children are in the second and third columns of Panel A, with the fourth column showing the differences in outcomes and background characteristics between the two cohorts. Similarly, the following three columns in Panel A detail summary statistics for the comparison group, while the last column reports the difference-in-differences between parents of children of different age groups. The table also includes the significance level for the difference-in-difference column. As shown, the mothers of two-year-olds reduce their level of full-time employment by close to four percentage points between 1997 and 2000. However, the mothers of five-year-olds somewhat increase their full-time attachment in the same period. The earnings difference also clearly reflects the withdrawal of mothers of two-year-olds from the post-reform labor market. However, for fathers, there is no indication of different outcomes depending on the age of the child, even though the same dip discussed earlier is apparent for the full-time attachment variable.

To explore further the credibility of the identifying assumption on common trends, we start by considering the child background characteristics in Panel B in Table 1. We note that there appears to be an increase in the number of families with two children in the post-reform comparison group. Turning to Panel C and parents' background characteristics, levels appear similar for the two groups of parents of two- and five-year-olds, with just a few exceptions. For example, we can see that parents of children from later cohorts are more likely to have finished college than are parents of older children. This possibly relates to a general education trend in the population during this period.[24] Lastly, there is a barely significant decrease in the share of parents living in densely populated areas. This may influence labor force attachment if there is a higher unemployment rate in rural areas. Remembering that I include a measure of municipality unemployment rates as a covariate, this should be accounted for in the regressions. To account for possible observable changes in the composition of the groups, I include covariates in all of the regressions (unless otherwise noted). I also add the covariates sequentially to investigate whether the slightly changing trends observed affect the results.

[24] Statistikkbanken, Statistics Norway.

Table 1: Summary statistics

	Treat/ pre	Treat/ post	D	Comp./ Pre	Comp./ post	D	D-in-D
Panel A: Outcome variable	Born 1995	Born 1998		Born 1992	Born 1995		
Mother full-time age 2/5	0.312	0.270	–0.042**	0.334	0.343	0.009*	–0.051**
Mother's earnings 2/5	122 858	116 190	–6 668**	136 785	142 523	5 738**	–12 406**
Father full-time age 2/5	0.726	0.649	–0.077**	0.730	0.653	–0.077**	0.000
Father's earnings 2/5	282 594	280 297	–2 297	294 533	294 788	255	–2 551
Panel B: Child Characteristics							
2 children	0.463	0.462	–0.001	0.455	0.463	0.009*	–0.010+
3 children	0.238	0.236	–0.002	0.240	0.238	–0.002	0.000
4 children	0.050	0.049	–0.001	0.057	0.050	–0.007**	0.006*
5 children or more	0.016	0.015	–0.001	0.017	0.016	–0.001	0.000
Sex=female	0.492	0.489	–0.003	0.487	0.492	0.005	–0.009
Panel C:							
M prior earn.	145 806	145 438	–369	120 884	122 858	1 974*	–2 342+
M minor part-t. prior	0.122	0.128	0.006+	0.165	0.166	0.001	0.005
M part-t. prior	0.116	0.121	0.005+	0.132	0.139	0.006*	–0.001
M full-t. prior	0.420	0.441	0.021**	0.299	0.334	0.035**	–0.014*
M high sch.	0.548	0.606	0.058**	0.492	0.556	0.064**	–0.006
M college	0.273	0.311	0.038**	0.253	0.281	0.028**	0.009+
M age	30.946	31.410	0.464**	30.560	30.946	0.386**	0.078
M immigrant	0.082	0.090	0.009**	0.078	0.087	0.009**	–0.001
M on welfare	0.029	0.032	0.003*	0.041	0.040	0.000	0.003
M urban area	0.741	0.743	0.002	0.729	0.740	0.010**	–0.009
F prior earn.	265 396	264 384	–1 012	284 867	282 594	–2 273	1 261
F minor part-t. prior	0.022	0.027	0.005**	0.016	0.020	0.004**	0.001
F part-t. prior	0.018	0.017	–0.001	0.016	0.017	0.002	–0.002
F full-t. prior	0.722	0.751	0.029**	0.732	0.759	0.028**	0.002
F high sch.	0.593	0.640	0.047**	0.579	0.617	0.039**	0.008
F college	0.258	0.278	0.021**	0.263	0.265	0.002	0.018**
F age	33.807	34.197	0.390**	33.414	33.807	0.394**	–0.004
F immigrant	0.076	0.084	0.008**	0.074	0.080	0.006*	0.002
F on welfare	0.038	0.032	–0.006**	0.044	0.039	–0.005**	–0.001
F urban area	0.738	0.741	0.003	0.728	0.738	0.011*	–0.007
Unempl.	0.027	0.023	–0.005**	0.042	0.027	–0.015**	0.010**
N	25,557	25,253		25,302	25,557		

Note: Mean or share of indicated variables with differences. Earnings are inflation adjusted with 1997 as the base year (in NOK). +, * and ** denote significance at the 10, 5 and 1 percent levels (two-sided t-test), respectively, and are reported for the difference-in-differences estimates.

6.2 Parental labor supply

Table 2 provides OLS estimates of the difference-in-differences coefficients obtained from the estimation of Equation (2), stepwise adding sibling, mother and father characteristics in Models 2, 3 and 4, respectively. In Model 5, I include municipality unemployment, while an interaction term between unemployment and the age of the child is included in Model 6. The first column provides the estimated results without the covariates. We can see that mothers of two-year-old children decrease their full-time attachment to the labor market by 4.8 percentage points after the introduction of the Cash-for-Care program. The estimated effect remains significantly negative, but decreases slightly as I add the covariates. We can see in particular that adding mother characteristics somewhat reduces the effect on the estimate. Nevertheless, the small decline assures us that compositional changes do not seriously bias the estimates. In the remaining specifications, I employ Model 6 as the preferred model, as it includes all specified covariates.

It is worth noting that a decrease of four percentage points in the participation rate of women is substantial, especially given that the mean participation rate is 31.5 percent. In other words, the reduction of four percentage points implies that 13 percent of mothers working 30 hours or more pre-reform have reduced their labor supply to a level below this threshold compared with the reference group.

Panel B of Table 2 reports the effect of the Cash-for-Care subsidy on the employment of mothers. As shown, about 2.4 percentage points of mothers withdraw entirely from the labor market because of this subsidy. This suggests not only that the subsidy causes a greater proportion of mothers to reduce their work to a part-time position, but also that some mothers actually stop working entirely.

Table 2: Labor supply of mothers

	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6
Panel A: Full-time employment only						
	–0.048**	–0.047**	–0.042**	–0.042**	–0.042**	–0.040**
	(0.006)	(0.006)	(0.005)	(0.005)	(0.005)	(0.005)
N	101,669	101,669	101,669	101,669	101,669	101,669
R^2	0.004	0.022	0.285	0.287	0.292	0.292
Mean	0.315	0.315	0.315	0.315	0.315	0.315
Panel B: Full- or part-time employment						
	–0.028**	–0.027**	–0.024**	–0.024**	–0.024**	–0.024**
	(0.005)	(0.005)	(0.004)	(0.004)	(0.004)	(0.004)
N	101,669	101,669	101,669	101,669	101,669	101,669
R^2	0.003	0.018	0.264	0.268	0.294	0.294
Mean	0.834	0.834	0.834	0.834	0.834	0.834
Included covariates						
Sibling char		X	X	X	X	X
Mother char			X	X	X	X
Father char				X	X	X
Unemployment					X	X
Unemployment × age						X

Notes: +, * and ** denote significance at the 10, 5 and 1 percent level, respectively. The OLS estimations in Panel A are based on Equation (2), with the outcome being whether the mother has a full-time attachment to the labor market, while in Panel B the outcome is whether the mother has any attachment to the labor market. We follow the first cohort of fully treated children born in 1998 as two-year-olds in 2000. Model 1 excludes the covariates. In the following five models, I include sibling characteristics, mother characteristics, father characteristics, the municipality-specific unemployment rate in the year in which the treated children turn two and (in Model 6) the municipal unemployment rate interacted with the age of the child. Robust standard errors (in parentheses) are clustered on the child's mother and account for heteroscedasticity and non-independence of the residuals across maternal labor force participation observed at different points in time. All specifications include municipality fixed effects. Source: administrative registers, FD-Trygd.

We now turn to Table 3 to investigate whether the decrease in maternal labor force participation resulted in a change of fathers' labor supply. Panel A reports results from the model specified in Equation (2), with the outcome being the full-time attachment of fathers. As shown, there is no evidence of any impact on the paternal labor supply. These results are stable under sequential inclusion of the various sets of covariates, adding to the robustness of the result. Recalling the potential bias in results because of the introduction of the paternity quota discussed in Section 3, it is reassuring to see that adding the background characteristics of fathers in Model 4 does not change the estimates. This suggests that the paternity quota does not bias the results. In Panel B, we can see that there is clearly no evidence for a change in the likelihood of fathers having any attachment to the labor force. Once again, very little changes when the covariates are included.

Table 3: Labor supply of fathers

	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6
Panel A: Full-time employment						
	0.001	0.002	0.003	0.000	0.000	0.001
	(0.006)	(0.006)	(0.006)	(0.005)	(0.005)	(0.005)
N	101,669	101,669	101,669	101,669	101,669	101,669
R^2	0.007	0.014	0.055	0.292	0.314	0.314
Mean	0.689	0.689	0.689	0.689	0.689	0.689
Panel B: Full- or part-time employment						
	-0.000	-0.000	0.000	-0.000	-0.000	-0.003
	(0.003)	(0.003)	(0.003)	(0.003)	(0.002)	(0.002)
N	101,669	101,669	101,669	101,669	101,669	101,669
R^2	0.000	0.008	0.051	0.188	0.304	0.304
Mean	0.949	0.949	0.949	0.949	0.949	0.949
Included covariates						
Sibling char		X	X	X	X	X
Mother char			X	X	X	X
Father char				X	X	X
Unemployment					X	X
Unemployment × age						X

Notes: +, * and ** denote significance at the 10, 5 and 1 percent level, respectively. The OLS estimations in Panel A are for Equation (2), with the outcome being whether the father has a full-time attachment to the labor market, while in Panel B the outcome is whether the father has any attachment to the labor market. I follow the first cohort of fully treated children born in 1998 as two-year-olds in 2000. Model 1 excludes the covariates. In the following five models, I include sibling characteristics, mother characteristics, father characteristics, the municipality-specific unemployment rate in the year in which the treated children turn two and (in Model 6) the municipal unemployment rate interacted with the age of the child. Robust standard errors (in parentheses) cluster on the child's father and account for heteroscedasticity and non-independence of the residuals across paternal labor force participation observed at different points in time. All specifications include municipality fixed effects. Source: administrative registers, FD-Trygd.

Keeping in mind from Section 2 that the effects on fathers would likely depend on their spouse's previous earnings, we now consider Table 4. This table provides the results of the subsample analysis across maternal education background and prior income. As shown, mothers with a college degree reduce their full-time attachment to a similar extent at the margin as mothers without a college degree. A similar pattern is clear when considering the subsamples of mothers with baseline earnings below and above the median. The estimated coefficients in Model 3 and 4 of Panel A in Table 4 are not significantly different from each other.

Turning now to Panel B in Table 4, we investigate whether the mean across the entire sample of fathers reported in Table 3 may mask heterogeneous effects across subsamples of spouses with different background characteristics. Given that we did not observe differing effects for mothers, it does not come as a surprise that the effects for fathers are also very similar across the subsamples.

Table 4: Subsample analysis: Full-time employment

	Model 1	Model 2	Model 3	Model 4
	Mother finished college	Mother did not finish college	Mother's prior earnings below median	Mother's prior earning above median
Mothers				
Age 2	–0.037**	–0.041**	–0.030**	–0.042**
	(0.010)	(0.006)	(0.006)	(0.008)
N	28,409	73,260	50,871	50,798
R^2	0.259	0.281	0.094	0.217
Mean	0.452	0.262	0.123	0.507
Fathers				
Age 2	0.010	–0.002	–0.002	0.005
	(0.009)	(0.006)	(0.007)	(0.007)
N	28,409	73,260	50,871	50,798
R^2	0.290	0.322	0.327	0.289
Mean	0.720	0.678	0.646	0.733

Notes: +, * and ** denote significance at the 10, 5 and 1 percent level, respectively. Estimations are based on OLS estimates of Equation (2), with the outcome in Panel A (B) being whether the mother (father) has a full-time attachment to the labor market. Parents of the first cohort of fully treated children (born 1998) are evaluated when the child turns two, in 2000. Covariates described in Section 5 are included. Robust standard errors (in parentheses) are clustered on the child's mother (father) and account for heteroscedasticity and non-independence of the residuals across parental labor force participation observed at different points in time. All regressions include municipality fixed effects. Source: administrative registers, FD-Trygd.

Even if no effects on paternal full-time attachment are found, it may be that full-time working fathers worked extra hours to compensate for the loss of a spouse's income. To explore this possibility further, we turn our attention to Table 5, which reports the results on a more continuous measure of labor supply, namely, annual earnings. Panel A details the results on the effect of the subsidy on the earnings of mothers across the same subsamples as in Table 4. It is clear from Model 1 and 2 that the effect at the margin is similar for educated and other mothers. Mothers with baseline earnings below and above the median reduce their earnings to the same extent.

Panel B in Table 5 displays the results for paternal earnings. We see that the reduced labor force participation of mothers across all of the subsamples does not have any positive impact on paternal earnings. If anything, there appears to be suggestive evidence that fathers with spouses earning below the median in the baseline year actually also reduce their labor supply. This effect is small, however, and barely significant at the 10% level.

Table 5: Subsample analysis: Earnings

	Model 1	Model 2	Model 3	Model 4
	Mother finished college	Mother did not finish college	Mother's prior earnings below median	Mother's prior earning above median
Panel A: Mothers				
Age 2	–7,862**	–10,352**	–10,741**	–7,259**
	(1,652)	(881)	(1,057)	(1,146)
N	28,409	73,26	50,871	50,798
R^2	0.522	0.487	0.262	0.428
Mean	180 525	108 129	73 439	183 356
Panel B: Fathers				
Age 2	–4,217	–1,788	–3,503+	–1,398
	(3,000)	(1,602)	(2,091)	(1,816)
N	28,409	73,260	50,871	50,798
R^2	0.494	0.444	0.454	0.531
Mean	330 885	262 262	266 830	296 065

Notes: +, * and ** denote significance at the 10, 5 and 1 percent level, respectively. Estimations are based on OLS estimates of Equation (2), with the outcome being maternal linear earnings. Earnings are inflation adjusted with 1997 as the base year (in NOK), and censored at the 99^{th} percentile. We include parents of the first cohort of fully treated children born in 1998, evaluated when the child turns two, in 2000. The covariates described in Section 5 are included. Robust standard errors (in parentheses) cluster on the child's mother (Panel A) and father (Panel B) and account for heteroscedasticity and nonindependence of the residuals across parental labor force participation observed at different points in time. All regressions include municipality fixed effects. Source: administrative registers, FD-Trygd.

6.3 Robustness

In order to explore further whether it is reasonable to expect similar trends across parents with children of different ages, the first two models in Table 6 provide estimates from a placebo model in which we assume that the reform took place three years before the actual treatment. If preexisting trends affected labor supply differently in the treatment and comparison groups, we would discern a treatment effect in the years before the Cash-for-Care subsidy was introduced. It is clear from column three in Table 6 that there is no indication of a preexisting diverging trend. This holds for both the mothers (Panel A) and fathers (Panel B).

Another concern raised in Section 4 was that the sample selection criteria could bias the results. Models 3 and 4 display the results from a specification that includes all parents, not just intact families. We can see that even without the earlier sample selection restriction, the results for both the full-time attachment and earnings of mothers (Panel A) and fathers (Panel B) are quite consistent. We do not confirm the small positive effect on paternal full-time attachment of increased earnings, and thus there should be no reason to alter the earlier conclusion concerning the main specification on this basis.

Table 6: Robustness

	Full time attachment Placebo	**Earnings Placebo**	**Full time attachment Unrestricted sample**	**Earnings Unrestricted sample**
Mothers				
	0.001	1,820	-0.029**	-8,075**
	(0.007)	(1,434)	(0.003)	(528)
N	99,761	99,761	239,735	239,735
R^2	0.002	0.006	0.252	0.521
Mean	0.309	127 991	0.290	120 537
Fathers				
	0.005	-185	0.010**	-4,820
	(0.007)	(2,280)	(0.003)	(3,559)
N	99,761	99,761	239,735	239,735
R^2	0.001	0.002	0.295	0.115
Mean	0.718	283 980	0.652	269 006

Notes: +, * and ** denote significance at the 10, 5 and 1 percent level. Estimations are based on OLS on equation (2) with outcome being full time attachment and linear earnings. Earnings are inflation adjusted with 1997 as base year (in NOK), and censored at the 99th percentile. Due to data restrictions model 1 and 2 does not include covariates or municipality fixed effects. The placebo sample is based on the same selection criteria as the main sample and is the same age, but belong to non-treated cohorts (born 1989, 1992 and 1995).The full sample includes the entire cohort of two and five year-olds. Model 3 and 4 include the covariates listed in table 1. In all models robust standard errors (in parenthesis) are clustered on the child's mother (panel A) and father (panel B) and account for heteroscedasticity and non-independence of residuals across parents' labor force participation observed at different points in time. Source: Administrative registers: FD Trygd.

7 Conclusion

There is increasing recognition that fathers represent an important family resource, beyond merely securing household income. The modern family needs the father at home because children benefit from his presence. Understanding better how extensive home care policies affect the division of labor in the household is consequently of great importance. This paper explores the variation caused by the introduction of a Cash-for-Care subsidy in Norway in 1998. This subsidy altered the relative productivities of market and home production. The principal focus has been to determine whether the withdrawal of mothers from the labor market because of this subsidy induces fathers to work more, as the seminal theoretical arguments in the work of Gary Becker imply.

The analysis utilized a comprehensive, longitudinal register database containing annual records for every person in Norway. I estimate difference-in-differences models by exploiting differences in the exposure of individuals to the program among families with similar structures. The main findings are that mothers reduced their attachment to the labor market substantially because of the reform. In fact, the decrease in the number of full-time working mothers is approximately four percentage points, implying that 13 percent of previously full-time working mothers began working part time.

However, I find no effect on fathers' degree of attachment to the labor market, and few effects, if any, on fathers' earnings. I explore several possible offsetting effects depending on the characteristics of their spouses, but these results persist. Put simply, fathers did not change their attachment to the labor market after the introduction of the subsidy, despite its large effects on mothers.

To the extent that specialization defines how the labor supply of spouses changes relative to each other, it may be that Becker's predictions hold to the extent that a relative increase in the productivity of home production induces one spouse to specialize at home. However, Becker's formal models also predict that the withdrawal of one spouse should lead the other to work more. I find no evidence for this latter prediction when considering the variation generated by the introduction of the Cash-for-Care subsidy. This might be partly explained by the substantial sum of money the subsidy delivered, possibly fully compensating for mothers reduced earning in many families.

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APPENDIX

Expected hours worked: Documentation on the variable

The variable 'expected hours worked' describes whether an individual works 4–19 hours, 20–29 hours or more than 30 hours per week. In 1999 and 2000, there was a significant increase in missing observations for this variable. There are two parallel explanations for this. First, from 1999 onward, companies were supposed to report exact hours worked to the employer register (Arbeidstakerregisteret). Previously, the companies would report values 1–3 for short and long part-time and full-time, respectively. In a transitional phase, many companies may have reported according to the earlier requirements, or not reported at all (this could lead to missing values). This may have caused many observations to be registered with the values 0 1 00, 0 2 00 or 0 3 00 for the exact hours worked. This could also explain the increase in missing observations.[25] Second, there was a change of tax scheme in 2000 that made it more profitable to remain self-employed (Alstadsæter and Thoresen 2008). This would affect the variable that measures attachment to the labor market, as we do have information on the degree of attachment for the self-employed.

To ensure that neither the change in registration routines nor the change in self-employment conditions affect our results, we report below the main results concerning the full-time attachment of parents, with an outcome variable that includes self-employed people. In this instance, we define full-time employment registration as having an employment code and earnings above 3G. While the full-time variable used in the main specification is simply employers' report of employees' status by the end of the year, the current specification will depend on the income ceiling. We specify 3G as the income ceiling to define full-time employment because it is approximately the lowest possible earnings one can receive for a full-time position according to the agreed official wage levels in 1997.[26] This implies that we will also include (exclude) some parents who work part (full) time and have high

[25] According to Jørn Ivar Hamre, Statistics Norway.
[26] St.prp. nr. 76 (1996–97) Om lønnsregulering for de offentlige tjenestemenn m.fl.

(low) wages. Table 7 reports the results from the regressions of Equation (2) with the covariates listed in Table 1 and with the revised definition of full-time attachment. We see that, if anything, the full-time variable found in the main specification underestimates the effect of labor force attachment for mothers. For fathers, the findings are generally consistent with those in the main body of the text.

Table 7: Main results for the labor supply of mothers and fathers

	Change in full-time attachment of mothers	Change in full-time attachment of fathers
Age2xyear2000	–0.065**	–0.004
	(0.007)	(0.004)
N	101,669	101,669
R^2	0.012	0.000
Mean	0.512	0.893

Notes: +, * and ** denote significance at the 10, 5 and 1 percent level, respectively. Estimations are OLS estimates of Equation (2), with the outcome being whether the mother/father had a full-time attachment to the labor market. We evaluate the parents of the first cohort of fully treated children (born 1998) when the child turns two, in 2000. The covariates in Table 1 are included. Robust standard errors (in parentheses) cluster on the child's mother (father) and account for heteroscedasticity and nonindependence of the residuals across parental labor force participation observed at different points in time. All regressions include municipality fixed effects Source: administrative registers, FD-Trygd.