Nordic Journal of Political Economy

Volume 40 | 2015 | Article 2

Crowding out Dad? The Effect of a Cash-for-Care Subsidy on Family Time Allocation*

Nina Drange[a]

[a]*Nina Drange,* Statistics Norway. E-mail: nina.drange@ssb.no

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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 variation 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.

JEL Codes: J13, J22, J24

***Acknowledgments:** The author is grateful to seminar participants and conference delegates for helpful comments and discussions on an earlier version of this paper, with special thanks to Venke F. Haaland, Mari Rege, Marianne Simonsen, Ingeborg Foldøy Solli, Kjetil Telle, Mark Votruba and one anonymous referee. The author also gratefully acknowledges the financial support of the Norwegian Research Council (160965/V10).

1. Introduction

To assist parents in mitigating possible conflicts between work and family life, generous parental leave schemes are offered in many countries.[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 possible implication of the theory of specialization within the family (Becker 1965, 1985, 1991, Gronau 1973) is that if the mother allocates more time to home production, the father will increase his efforts in market production. Thus, if mothers reduce their labor supply, as is a typical intention of parental leave or home care schemes, and the mother's income is not fully compensated by the scheme, fathers may work more to compensate for the income loss. Consequently, parental leave policies and in particular schemes that offer only partial income replacement, could in some families have the unintended effect of causing the fathers to specialize further in labor market production.

The aim of this paper is to investigate how specialization patterns develop in families eligible for home care schemes. 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 between home and market production by both parents in two-parent families.[2] The level of subsidy involved in this scheme was substantial, and thus may have replaced lost earnings entirely in families where the mother had low earnings potential. It is likely that the behavior of fathers in these families differ from the one 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 predictions from theory (Becker 1965, 1985, 1991, Grounau 1973), we expect to observe relatively less specialization in families where the mother is well-educated, because a mother with education will likely face higher wages in market production. I investigate this empirically by comparing outcomes in a subsample of families where the mother has completed a college degree with outcomes of families where the mother has lower educational accomplishment.
A study from 2004 (Naz 2004) explores specialization in families with partly treated children around the introduction of the Cash-for-Care subsidy using survey data. Naz does not find support for an increase in fathers' labor supply, although the effect estimate is quite imprecise for this particular outcome. My paper adds to the literature by taking advantage of registry data of the full population of Norway, allowing both for a substantially larger sample size, as well as for a finer analysis based on pre-treatment characteristics. Moreover, I am able to isolate the effect of the subsidy on families with fully treated children.

[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).

[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).

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 any employment and full-time employment, 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 mothers' full-time employment by about four percentage points. Interestingly, mothers with low and high education responded similarly to the policy change at the margin. The reduction of mothers' full-time employment is consistent across groups with different levels of education, and for different levels of former earnings. 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. This holds regardless of their spouse's former earnings.

The remainder of the paper is structured as follows. Section 2 presents predictions from a model of the intra family allocation of time (Gronau 1973), and provides a brief overview of the empirical literature on specialization in the family. 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 Theory and empirical evidence

2.1 The intra family allocation of time

Most families with children divide their time between working at home taking care of the children, cooking and cleaning, and spending time in paid labor. Drawing on the work of Gary Becker (1965, 1985, 1991) and others, Gronau (1973) formulates a more general theory of the intra family allocation of time. To help us think about the possible effects of the Cash-for-Care subsidy in a formal specialization framework, I will lean on the implications of Gronau's model.[3]

A household's utility depends on each of its two members' leisure time (L), and the household consumption of market (M) and home (H) goods: $U = U(M, H, L_1, L_2)$. Gronau shows that even a small initial difference in the wage rate and/or in the productivity in home production, will lead to specialization among household members. It is the household member with the lowest wage rate (or the highest productivity in household production) who allocates his/her labor supply between the household and the market. Home goods are produced using time and market inputs. The optimum allocation of labor between home and work hours in the market will be where the value of home time equals the market wage of the lowest paid household member, and will also depend on the wage rate of the other household member.

Furthermore, Gronau shows that an increase in the wage of the highest paid household member will be accompanied by an increase in the amount of time the lowest paid household member spends at home, and consequently by a decline in the amount of time he/she devotes to the market sector. If the wage rate of the lowest

[3] This model is carefully spelled out in Gronau (1973). To be able to discuss the introduction of a subsidy on home care I will have to make some assumptions about the subsidy. These assumptions will be made explicit throughout the text.

paid household member increases, the leisure of the highest paid household member increases, and his/her labor supply declines.

Turning now to the Norwegian setting, we start by noting that despite a substantial increase in female labor force participation since the 1970s, fathers still work more than mothers.[4] Thus, in the following I will assume that it is the mother who earns the lowest market wage and divides her time between home and market production, while the father works in market production. This is clearly not the case in all Norwegian families, but it simplifies the discussion of the subsidy and its effects using the theoretical framework provided in Gronau's model.

We consider the introduction of the Cash-for-Care subsidy as an incentive to produce more child care in the home, i.e. an incentive to increase home production, in families with eligible children. The implicit market wage rate of the spouse dividing her time between market and home production will fall, because buying center-based child care now imply not only the fee for a child care slot, but also a foregone subsidy. Leaning on the aforementioned model of the division of labor in the household, the implication of such a subsidy on home production will be that the mother[5] works more in the home.

How the family allocates the time of the father will depend on the size of the subsidy, as well as on how much time the mother allocated to market work (depending on the relative wage rate of the mother and father) to optimize family utility before the subsidy was introduced. I will elaborate on three different predictions[6]:

If the subsidy is large enough to exactly compensate for the mother's foregone earnings in the labor market, the family-income is unchanged and the family's incentives for paternal work in the labor market remain unchanged.

If the optimum family allocation of labor before the subsidy was introduced was such that only the father worked in the market, a cash subsidy will be like a lump sum transfer. Gronau (1973) shows that an increase in other sources of income will result in increased leisure by both the market and the home working spouse. In this case the market working spouse will work less.

Lastly, if the mother worked full time before the introduction of the subsidy, the family may have a lower income than before. The mother is now at home taking care of the toddler, and the subsidy might not be large enough to compensate for this. In this case, the family may be better off by allocating more of the father's time to market production to maintain the earlier level of income.

Theory predicts increased labor supply for the father only in the case where the mother had a strong labor force attachment prior to the introduction of the subsidy. I will investigate if this holds by comparing labor supply of fathers whose spouses earned below and above median income respectively, as well as labor supply of fathers with spouses having low vs. high education.

[4] When considering summary statistics in Table 1 we see that income the year before birth for mothers and fathers with a child born 1998, is 145 000 and 245 000 respectively. The corresponding share of mothers and fathers working full time is 44 % and 75 %.

[5] 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.

[6] Note that I to simplify assume that buying child care is a 0-1 decision: Either you hold a full time slot, receiving no subsidy, or you receive the subsidy and do not hold a child care slot. Furthermore, I assume that when the mother is at home with the toddler, she does not increase household production beyond looking after the child.

2.2 Empirical evidence

The existing empirical literature on specialization effects of parental leave or home care subsidies is sparse[7], 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,[8] it is only because they spend less time on leisure activities, not because they work less.

A study by Lundberg and Rose (2000) employs fixed- and random-effects models in order to explore changes in specialization occurring after the birth of a couples' first child. Comparing parents and non-parents, they find significant differences in the labor supply of the two groups even before the parents-to-become actually have a child. Women who become mothers earn less than other women before birth. Furthermore, Lundberg and Rose find that a child is born, the father works more and the mother works less if the mother has had a career break connected to the birth. Moreover, the mothers earn less per hour and the fathers earn more. When the mother has not had any career break due to the birth, she subsequently works fewer hours, but her wage is not reduced. Men with spouses that do not choose to have a career break work fewer hours but earn more per hour.

While a number of studies have also investigated the effects of parental leave and home care schemes on maternal labor supply, few have focused on the possible specialization effects.[9] As mentioned in the introduction, 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 production, Naz (2004) found that the subsidy decreased the labor supply of mothers, but did not affect fathers. The present analysis differs from Naz (2004) in several respects. First, by employing registry 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.[10] Nevertheless, despite these and several other key differences, my findings are largely consistent with Naz (2004).

[7] 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).

[8] 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.

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

[10] Two- and three-year-olds in 1999 were partly treated as one- and two-year-olds in 1998. See Section 3 for details.

3 Institutional setting

3.1 The Cash-for-Care program

After the election for the Norwegian parliament in September 1997, an alliance consisting of the Christian Democratic party, the Centre party and the Liberal party formed a new government. One of the issues on their political agenda was the introduction of a Cash-for-Care subsidy. The cash-for-care Act was passed in the parliament in June 1998 (Norwegian Ministry for Children and the Family (1998)).

The new 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),[11] 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 (2012) 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 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.[12] 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

[11] The transfer was tax-free.

[12] 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.

subsidy expired. This is because if the Cash-for-Care subsidy had a persistent effect on 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.[13] It is also worth noting that very few fathers received the subsidy.

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

	Recipients All	Fathers	Mothers	Children 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

[13] 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.

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[14] (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,[15] 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 full wage compensation or, alternatively, 52 weeks with 80 percent wage compensation,[16] 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 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

[14] OECD Family Database, based on statistics from 2007.
[15] OECD Labor Market Statistics (http://stats.oecd.org) and Statistics Norway (1998).
[16] In 2009, parental leave extended to 46 weeks of full compensation or 56 weeks of 80 percent compensation.

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,[17] 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,[18] 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.[19] 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 potential exogenous variation in framework conditions facing parents of one- and two-year-old children before and after the policy reform. The 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:

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

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).[20] The difference-in-

[17] Gelbach (2002) found that the maternal labor supply increases when the mother's oldest child starts school.
[18] See, for instance, Baker et al. (2008).
[19] Statistics Norway (2003).
[20] Alternatively, I could use a triple difference approach as in Schøne (2004), adding a third difference measuring the change in labor force participation for the same individual from a pre- to a post-period. Given that I include background characteristics that control for former labor force participation, my approach should give similar estimates. The advantage of the double difference is that is allows me to control for a rich set of parental characteristics at baseline.

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:

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

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.

Increased labor supply In particular, 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 I investigate closely if trends seem to differ pre-reform.

Fertility Policies implemented to mitigate costs connected to the birth of a new child might increase fertility (see for instance Lalive and Zweimüller (2009)). A sample selection criteria for all four groups in this study (comparison and treatment group before and after the policy change) is that the children should not have a new sibling the year they turn five years old. A possible fertility effect of the subsidy could affect the groups differently trough this sample selection criteria. The youngest cohorts have younger mothers when the policy is introduced, and are thus more likely to have new siblings if fertility increases. In fact, the cohort born 1992 have already turned five when the subsidy is introduced, so in this group selection out due to the subsidy is not allowed by definition. I address this concern in the following manner: I construct a sample of mothers with children born in the relevant years *without* imposing any selection criteria. I proceed to run the exact same regression as in the main specification, with the exact same covariates, but with the outcome being if the mother has a new baby by year five. If the subsidy indeed increased fertility to a larger extent for the youngest cohorts, this specification should produce a positive coefficient when implemented on the unrestricted sample.

Timing of fertility An important prerequisite for the empirical strategy to be valid is that there should be no selection into the treatment group. If parents could anticipate the introduction of the subsidy and time the birth of their child so that the child became eligible, this could bias the results. Keeping in mind from Section 3.1 that the final decision on the implementation of the subsidy came in June 1998, timing of birth in 1998 should not be of major concern. There was, admittedly, a public debate on the issue prior to the election in September 1997. Some parents might have wanted to postpone conception to after the election when they would possess more information on whether the subsidy would actually be implemented. The result would be that some children that would otherwise have been born late 1997 were instead born in 1998. To explore possible fertility effects I will display the number of children born by month for the cohorts 1996 (benchmark), 1997 and 1998.

Concurrent programs 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 1st 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.

Marital stability 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 how this would affect the results. If increased marriage stability implies that a larger share of mothers with a full-time work potential is included in the post-reform treatment group, there will be a downward bias in the estimates. If, however, the mothers who remain married to a lesser extent have a full-time work potential, the estimate will be biased upward. To investigate if possible sample selection may affect the results, I run a robustness check on a sample without restrictions.

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 age five. 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 not employed.[21] 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.[22] 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.[23] We code all missing observations on earnings as zero.[24] I do not specify the log of earnings as is sometimes typical in this sort of

[21] 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).

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

[23] 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.

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

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 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[25], the number of children (0, 1, 2, 3, 4, ≥5),[26] 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.

For the robustness checks I construct a sample identical to the main sample, but without applying the restrictions of excluding children with a younger sibling at age five or children with parents living apart at age five. To check for possible effects on fertility, I construct a dummy equal to one if the mother of a child in the sample gives birth to a new child by year five, and zero if she does not.

6 Empirical results

6.1 Summary statistics

Figure 3 depicts the trends in the share of full-time employed mothers and fathers of children in the age groups included in the analysis.[27] 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- 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.

[25] 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.

[26] Parenthetical documentation on any control variable indicates the ranges of the series of categorical variables that characterize the specific trait.

[27] 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.

Turning to the details for fathers, in Panel B of Figure 3 it appears that the trends in full-time employment 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 employment

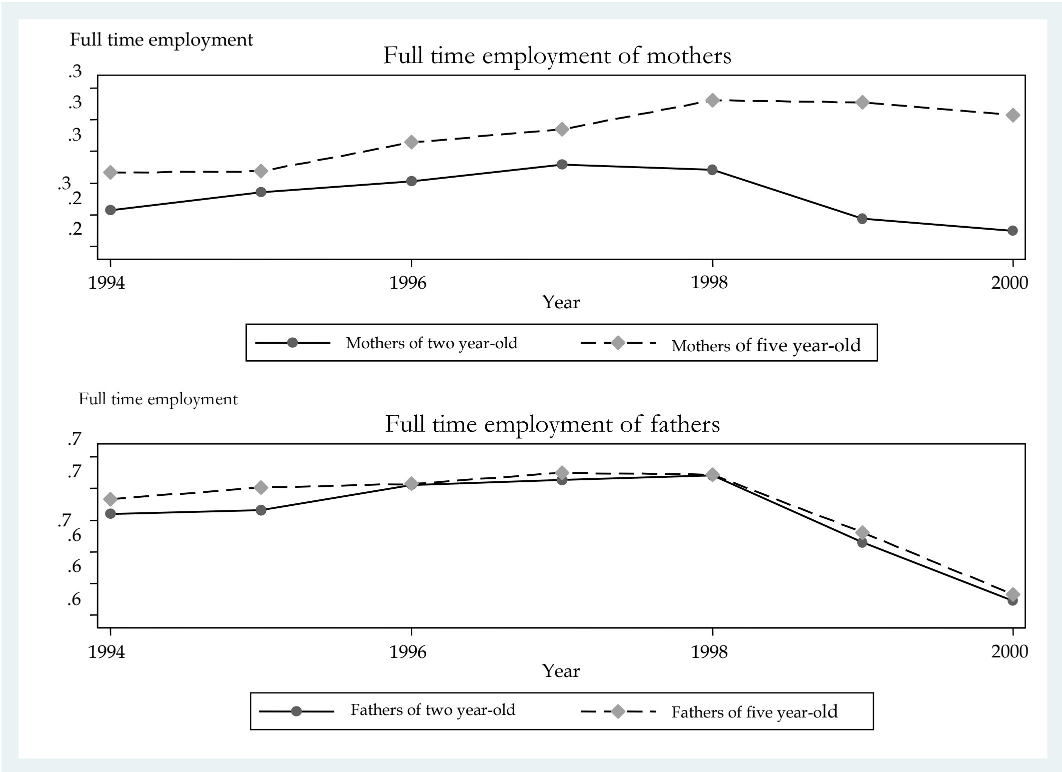


Figure 4 provides summary statistics on the earnings of mothers and fathers. As shown, from 1998 onward, the trends in the earnings of mothers of two- and five-year-olds begin to diverge, in line with what we have already observed concerning the likelihood of full-time work in Figure 3. The earnings 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

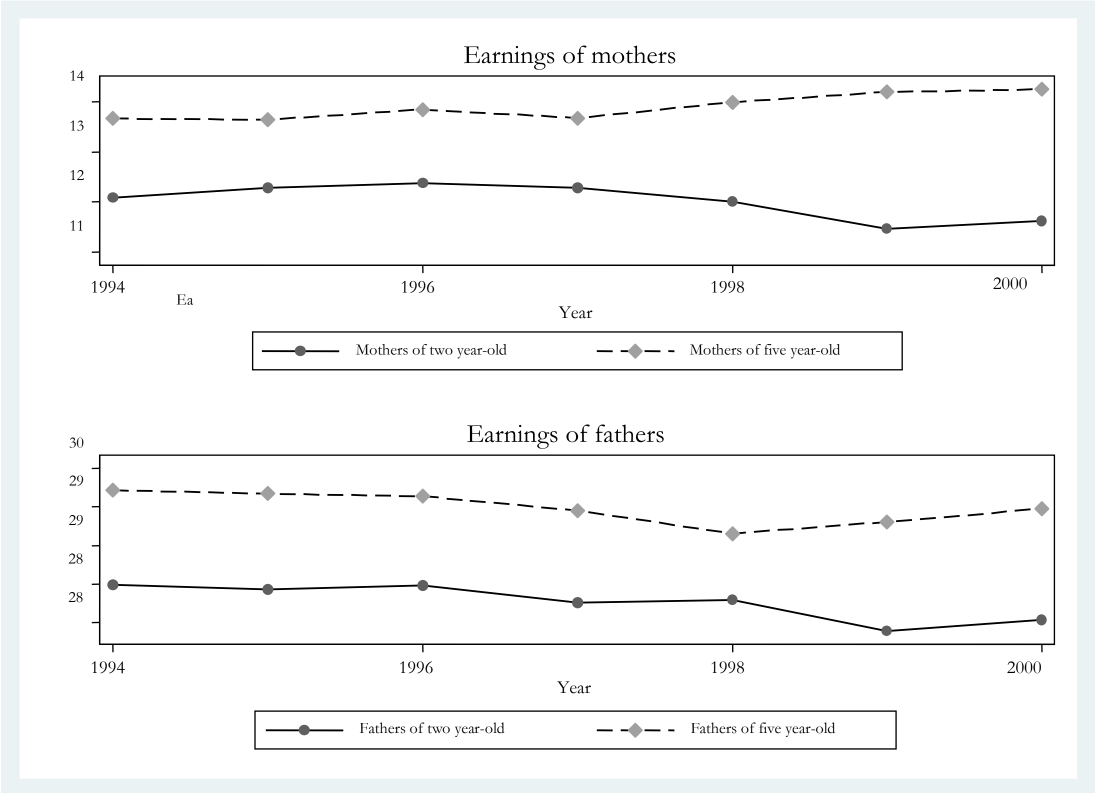


Table 1 details similar summary statistics. The outcomes for the mothers and fathers of treated children are in the second and third columns of Panel A, with the fourth column showing the differences in outcomes and background characteristics between the two cohorts. Similarly, the following three columns 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 employment 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 employment 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.[28] 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

[28] Statistikkbanken, Statistics Norway.

a higher unemployment rate in rural areas. Since 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.

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 employment 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 is full time employed, 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 is mirrored by a change of fathers' labor supply. Panel A reports results from the model specified in Equation (2), with the outcome being the full-time employment of fathers. As shown, there is no evidence of any impact on the paternal labor supply: The effect estimates are very small and precisely estimated. 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 no evidence for a change in the likelihood of fathers being employed. 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 is full-time employed, 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 employment 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. We see that there is little indication that fathers' respond differently to the subsidy depending on their spouses' former earnings or education.

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) is full-time employed. 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 significant effects on paternal full-time employment are found, it may be that full-time working fathers start working extra hours to compensate for the loss of their spouses' 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, 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. Again, we remember the predictions from Section 2 that fathers having a spouse with low earnings could possibly work less (or work the same) after the subsidy was introduced, whereas fathers with high earning spouses might work more. However, we see that the reduced labor force participation of mothers with former earnings above the median is not mirrored by an increase in paternal earnings. There appears to be suggestive evidence that fathers with spouses earning below the median in the baseline year reduce their labor supply in line with the predictions of the model. This effect is small, however, and barely significant at the 10% level.

Lundberg and Rose (1999) note that there are two important dimensions to the issue of household responses to the birth of a new child. In addition to the traditional measure of specialization, defined as the difference between the work hours or earnings of the husband and wife, they also suggest that market intensity is of separate interest. Market intensity is the sum of work hours or earnings. From Table 5, we see that there is clearly a decrease in market intensity of couples across all groups caused by the introduction of the Cash-for-Care subsidy.

Table 5: Subsample analysis: Earnings

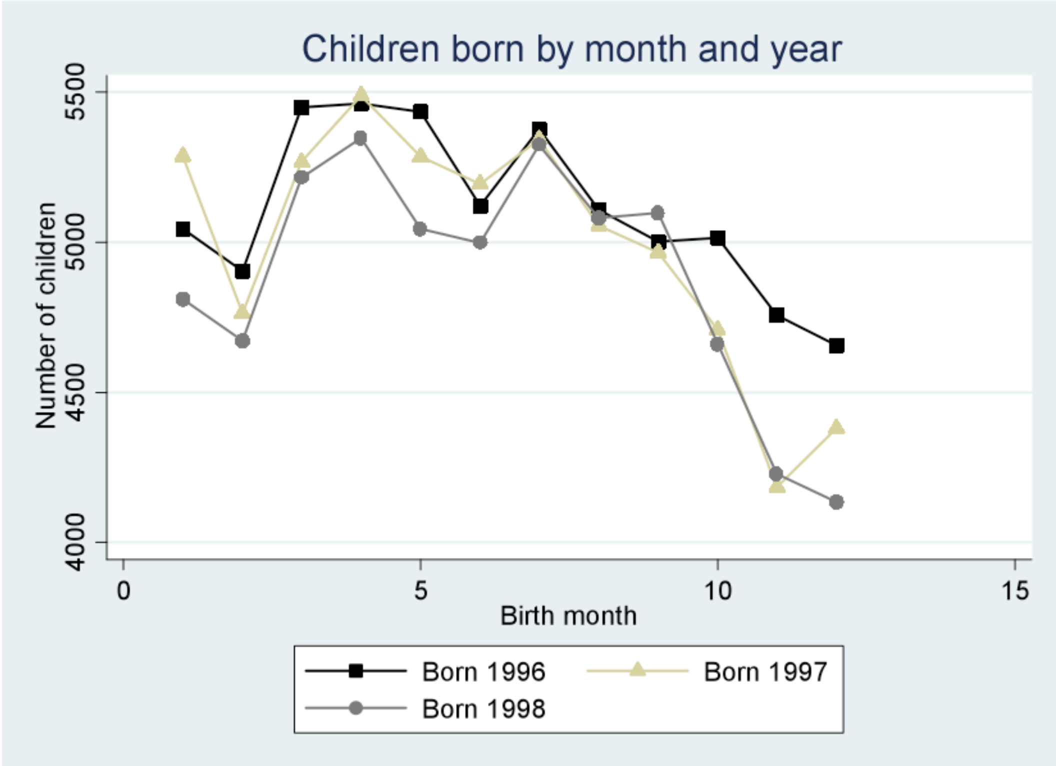
	Model 1	Model 2	Model 3	Model 4
	Mother finished college	Mother did not finish college	Mother's prior earnings below median	Mother's prior earning above median
Panel A: Mothers				
Age 2	–7,862**	–10,352**	–10,741**	–7,259**
	(1,652)	(881)	(1,057)	(1,146)
N	28,409	73,26	50,871	50,798
R^2	0.522	0.487	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 99th 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

We start by considering possible selection into the treatment group, i.e. into being born 1998. Keeping in mind that the issue of introducing a Cash-for-Care subsidy was on the political agenda during the election campaign prior to the election in September 1997, we could for instance imagine that parents that would otherwise wanted to conceive a child early 1997, would wait until after the election in order to have greater certainty about the subsidy. This would result in a reduction of children born late 1997, and a surge in children born second half of 1998. Figure 5 below displays the number of children born by month in the years 1996, 1997 and 1998. Little indicates that this is the case, there are small variations throughout the year, but nothing stands out.

Figure 5: Children born by year and month



To further explore 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 the first column 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. We are in particular worried about a possible subsidy effect on fertility. Model 3 in Table 6 reports results from a regression on whether the child has a new sibling by year 5. It is reassuring that the coefficient is a fairly precisely estimated zero, suggesting that the cohorts included in treatment and comparison groups are not experiencing different fertility trends by year five due to the subsidy.

Models 4 and 5 display the results from a specification that include all parents in a sample without restrictions. We can see that even without the earlier sample selection restriction, the results for both the full-time employment 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 employment on increased earnings, and thus there should be no reason to alter the earlier conclusion concerning the main specification on this basis. Little suggest that the main findings are driven by sample selection.

Table 6: Robustness

	Placebo		Fertility		
	Model 1	**Model 2**	**Model 3**	**Model 4**	**Model 5**
	Full time attachment	**Earnings**	**Effect on fertility by age 5**	**Full time employment Unrestricted sample**	**Earnings Unrestricted sample**
Mothers					
	0.001	1,820	-0.003	-0.029**	-8,121**
	(0.007)	(1,434)	(0.004)	(0.003)	(528)
N	99,761	99,761	239,701	239,701	239,701
R^2	0.002	0.006	0.37	0.252	0.520
Mean	0.309	127 991	0.429	0.290	120 527
Fathers					
	0.005	-185	- - -	0.008**	-4,861
	(0.007)	(2,280)	- - -	(0.003)	(3,549)
N	99,761	99,761	- - -	239,701	239,701
R^2	0.001	0.002	- - -	0.291	0.114
Mean	0.718	283 980	- - -	0.652	269 142

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 employment 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, 4 and 5 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. 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 explore if the subsidy induced families to allocate more of the mother's work hours to the home and more of the father's work hours in market production, as the seminal theoretical arguments in the work of Gary Becker imply.

The 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 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 of full-time employed mothers is approximately four percentage points, implying that 13 percent of previously full-time employed mothers began working part time.

Theory suggests that the mother's earning opportunities in market production will affect the degree of specialization within the family (Gronau 1973). In families where the mother had high earnings before the subsidy was introduced, we might expect the father to work more. In families where the mother had low earnings, we should expect unchanged labor supply by the father, and in some cases a decrease. Although mothers with both former high and low earnings reduced labor supply substantially after the subsidy was introduced, I find no significant effect on fathers' degree of attachment to the labor market, and small and barely significant 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 is defined by how spouses' labor supply changes relative to each other, the predictions from theory hold. A relative increase in the value of home production induces one spouse to specialize at home. However, theory also predicts that the household would allocate more of the fathers' time to market production, particularly in families where the mother initially had high earnings. 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.

References

Alstadsæter, A. and T. O. Thoresen (2010): Shifts in Organizational Form under a Dual Income Tax System, *FinanzArchiv/Public Finance Analysis*, 66(4), 384-418.

Angrist, J. and W. Evans (1998): Evidence from Exogenous Variation in Family Size. *American Economic Review*, Vol. 88, No. 3, 450–477.

Baker, M., J. Gruber and K. Milligan (2008): Universal Child Care, Maternal Labor Supply, and Family Well-Being. *Journal of Political Economy*, Vol. 116, No. 4, 709–745.

Becker, G. S. (1965): A Theory of the Allocation of Time, *Economic Journal*, Vol. 75, No. 299, 493–517. (Reprinted in Becker [1976].)

Becker, G. S. (1985): Human Capital, Effort, and the Sexual Division of Labor, *Journal of Labor Economics*, Vol. 3, No. 1, Part 2, S33–S58.

Becker, G. S. (1991): *A Treatise on the Family*. Harvard University Press. Enlarged edition.

Bettinger, E., Hægeland, T. and M. Rege (2014): Home with Mom: The Effects of Stay-at-Home Parents on Children's Long-Run Educational Outcomes, Journal of Labor Economics, 32(3), 443 - 467.

Cools, S., J. H. Fiva and L. J. Kirkebøen (2015): 'Causal Effects of Paternity Leave on Children and Parents', Scandinavian Journal of Economics, 117, 801-828. Available online: http://onlinelibrary.wiley.com/doi/10.1111/sjoe.12113/abstract

Drange, N and M. Rege (2013): Trapped at Home: The Effect of Mothers' Temporary Labor Market Exits on Their Subsequent Work Career, Labour Economics, 24, 125-136.

Gelbach, J. B. (2002): Public Schooling for Young Children and Maternal Labor Supply. *American Economic Review*, 92(1): 307–322.

Gronau, R. (1973): The Intrafamily Allocation of Time: The Value of the Housewives' Time. *The American Economic Review*, 63(4), 634-651.

Hardoy, I. and P. Schøne (2008): Subsidizing "Stayers"? Effects of a Norwegian Child Care Reform on Marital Stability. *Journal of Marriage and Family*, 70(3), 571-584.

Harris, K. M., F. F. Furstenberg and J. K. Marmer (1998): Paternal Involvement with Adolescents in Intact Families: The Influence of Fathers over the Life Course. *Demography*, 35(2), 201-216.

Kjeldstad, R. and M. Rønsen (2002): Enslige foreldre på arbeidsmarkedet 1980–1999. En sammenligning med gifte mødre og fedre. Statistical Analyses 49, Statistics Norway (in Norwegian).

Lalive, R. and J. Zweimüller (2009): How Does Parental Leave Affect Fertility and Return to Work? Evidence from Two Natural Experiments. *Quarterly Journal of Economics*, 24(3), 1363–1402.
Lamb, M., J. Pleck, E. Charnov and J. Levin (1987): A Biosocial Perspective on Paternal Behavior and Involvement. In J. Lancaster, J. Altman, A. Rossi and L. Sherrod (Eds.), *Parenting Across the Lifespan: Biosocial Perspectives* (11–42). New York: Aldine-De Gruyter.

Lundberg, S. and E. Rose (1999): The Determinants of Specialization Within Marriage. University of Washington, working paper.

Lundberg, S. and E. Rose (2000): Parenthood and the earnings of married men and women. *Labour Economics*, 7, 689-710.

Lundberg, S. and E. Rose (2002): The Effect of Sons and Daughters on Men's Labor Supply and Wages. *The Review of Economics and Statistics*, 84(2), 251-268.

Naz, G. (2004): The Impact of Cash-benefit Reform on Parents' Labor Force Participation. *Journal of Population Economics*, 17(2), 369–383.

Norwegian Ministry for Children and the Family (1998): Lov om kontantstøtte til småbarnsforeldre (The Cash-for-Care act), LOV 1998-06-26 no 41.

Norwegian Ministry for Children and the Family (1995): Pappa kom hjem. Utredning av utvalg oppnevnt av Barne- og familiedepartementet (in Norwegian), NOU 1995:27.

Norwegian Ministry for Planning and Coordination (1997): Om lønnsregulering for de offentlige tjenestemenn m.fl. St.prp. nr. 76 (1996–97)

OECD (2007): Babies and Bosses. Reconciling Work and Family Life. A Synthesis of Findings for OECD Countries. Report, OECD.

OECD (2011): OECD Family Database, OECD, Paris.

Rege, M. and I. F. Solli (2013): The impact of paternity leave on fathers' future earnings, Demography, 50(6), 2255-77.

Repo, K., T. Rissanen and J. Sipilä (2010): Cash-for-Childcare. The Consequences for *Caring Mothers*. Edward Elgar Publishing, UK.

Rønsen, M. (2001): Market Work, Child Care and the Division of Household Labor. Adaptations of Norwegian Mothers Before and After the Cash-for-Care Reform. Reports 2001/3, Statistics Norway.

Ludsteck, J. and U. Schönberg (2014): Expansions in Maternity Leave Coverage and Mothers' Labor Market Outcomes after Childbirth, Journal of Labor Economics, 32(3), 469-505

Schøne, P. (2004): Labor Supply Effects of a Cash-for-Care Subsidy. *Journal of Population Economics,* 17(4), 703–727.

Statistics Norway (1998): Kindergartens 1998. Official Statistics of Norway, C 583.

Statistics Norway (2003): Kindergartens 2003. Official Statistics of Norway, D 328.

Tamis-LeMonda, C. S. and N. Cabrer (Eds.) (2002): *Handbook of Father Involvement.* Mahwah, NJ: Erlbaum.

Tronstad, K. R. (red). (2007): Fordelingen av økonomiske ressurser mellom kvinner og menn. Inntekt, sysselsetting og tidsbruk. Report 2007/1, Statistics Norway (in Norwegian).

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.[29] 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 employment 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.[30] This implies that we will also include (exclude) some parents who work part (full) time and have high (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 employment. 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.

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

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

	Change in full-time employment of mothers	Change in full-time employment 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.