Balancing family and work: the effect of cash benefits for working mothers*

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Abstract

The aim of this paper is to measure the potential effect of a family policy introduced in Spain in 2003 that provides working mothers with a monthly cash benefit of 100 euros per child aged under 3 years. We explore the effect of the policy on eligible women's labour market participation. In the tradition of the policy evaluation literature we use a difference-in-differences-in-differences (DDD) estimation approach. Our results support a small but significant positive effect of the policy. We find that since the implementation of the policy the labour market participation rate for mothers of children aged under 3 has undergone an increase 2.926 percentage points higher than for non-policy-eligible females. This represents 5.164% of their average labour market participation in 2002, the year before the policy was implemented. We find that the effect is not homogeneous across educational groups, and, in particular, there is a stronger potential positive effect for women who have only completed elementary or high school education than for college graduated females.

JEL Codes: J13, J18, C51

Keywords: female participation, child care benefit, policy evaluation, difference-in-differences-in-differences estimation

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I. Introduction

At the Lisbon Summit held in June 2000 the EU governments announced that one of their primary goals was to reach a total employment rate of 70% by the year 2010. Policies aimed at increasing female labour participation of mothers are considered crucial to achieving this target. However, as is reported by Del Boca and Pascua (2002), substantial differences exist between southern countries, like Spain or Italy, and northern countries, like Norway or Sweeden, in terms of female participation in the labour market and in terms of family policies attempting to improve work incentives for parents.

In the case of Spain the evidence suggests that there are difficulties in reconciling female participation and motherhood. Firstly, the employment rate for mothers is lower in Spain than in most OECD countries. Gutiérrez-Domenech (2005) reports that Spain, together with West Germany, are the countries with by far the highest long-term decline in employed women after the first birth. Secondly, the use of formal child care arrangements for children aged under 3 years is very rare in Spain. As is reported in OECD (2001), the proportion of children aged under 3 years using formal child-care arrangements is only 5.0%, in sharp contrast to the average 24.7% for the European countries. Thirdly, the Spanish Labour Population Survey (2004) clearly indicates that the main reason why Spanish females do not participate in the labour market is related to family responsibilities. For the group of people aged under 45 years, only 0.6% of males report family responsibilities as the main reason for not participating in the labour market; however, this figure is 28.7% in the case of females. Fourthly, Spanish regulations concerning maternity leave establish 16 weeks of paid leave after birth, a shorter period than in most of the European countries where the average is about 25 weeks, according to OECD (2001). Finally, the fertility rate in Spain is one of the lowest in the EU-15 countries (Eurostat (2007)).

Encouraging mothers to participate in the labour market is a key mechanism to reduce the labour participation gender gap. On the one hand, the group corresponding to mothers of young children is a subset of the population with one of the lowest labour participation rates, see for example OECD (2002). On the other hand, spells out of the labour market for maternity reasons can erode human capital and reduce the accumulation of labour market experience, and thus result in a decrease of future potential wages and in a reduction of the incentives to join the labour market later in life.

Empirical and theoretical studies support the prediction that increasing child care cost lowers the female participation probability. Several studies provide evidence suggesting that the lower rate of participation shown by mothers of preschoolers in comparison to all other females is closely related to high child care costs faced by these females (see, for example, Blau and Robins (1988), Connelly (1992), Del Boca and Pascua (2002), Ferrero and Iza (2004), Attanasio et al. (2004)). At the same time there is a political debate regarding family policies trying to encourage mothers to enter the labour market. Recent measures adopted by the Spanish government represent a sound example of this kind of initiative. Specifically, in January 2003 the Spanish government passed a new family policy aimed at reconciling motherhood and labour market participation. The policy provides working mothers of children aged under 3 years with a monthly cash benefit of 100 euros per child.

The target of this paper is to measure the direct potential effect of this policy on the participation of mothers , i.e., its impact on the participation of mothers of children aged under 3 years, and thus, to contribute to the discussion about the convenience of increasing the amount of public resources devoted to the promotion of the participation of mothers in the labour market. We focus on this direct effect given the short period of time that has elapsed since the introduction of the policy. However, there may be a potential indirect effect on female participation of mothers of children aged over 3 years resulting from a stronger attachment to the labour market during the early maternity ages. Finally, the policy may also have effects on fertility decisions.

Through their impact on female labour market participation, the policy may also have effects on the income distribution across households (see for example Danziger (1980) and Cancian and Reed (1998) for the US and by Alba and Collado (1999) for Spain). Empirical evidence suggests a positive correlation between participation rates and educational attainment, as well as marital sorting by education, which lead us to think that the policy adopted in Spain can be non-innocuous for income distribution. Although this issue goes beyond the objectives of this paper, our findings may be relevant for such analysis.

Following the tradition of the policy evaluation literature this paper uses the 'difference-in-differences-in-differences' (hereafter DDD) estimation approach to measure the potential effect of the policy. The essence of this approach is to estimate the variation in the probability of participation of eligible women (in our case, mothers of children aged under 3 years) before and after the policy and compare it to the variation of the estimated participation probability of non-eligible women. First, we provide evidence of a sizeable growth rate differential in labour market participation between eligible and non-eligible women after the policy implementation. We then proceed to estimate an econometric model to perform the DDD approach. Based on our estimates, we claim that the potential effect of the cash benefit policy on the labour market participation of eligible females has been small but positive. Specifically, the increase in labour market participation for mothers of children aged under 3 years was, after the policy was implemented, 2.926 percentage points higher than the increase experienced by non-eligible females. This represents 5.164% of the before-policy average participation rate of the policy-eligible women. It is important to note that the effect of the policy is quite different across educational groups, it being stronger for mothers who completed elementary or high school education and weaker for college graduated mothers. In particular, for those mothers with elementary education the number is 8.690%, for those with high school education is 5.670%, and, for those with college education completed the figure is only 3.079%.

Our paper is related to others in the literature analyzing family policies. In particular, Schone (2004) and Naz (2004) examine the effect of a reform in Norway which provides mothers of children aged under 3 years with a cash benefit, irrespective of their working status. They find that the policy increases intra-household specialization of work and hence reduces the labour market participation of women. Francesconi and Van der Klaauw (2007) focus on the effect of 'in work' benefits to single mothers in the United Kingdom and find that the reform led to a substantial increase in female employment rates, of about five per cent. We contribute to this literature by providing new evidence on the labour market effects of a cash benefit for working mothers of young children during the period 2003-04 and new estimates of the elasticity of labour supply to childcare costs. Furthermore, there is an additional contribution derived from the fact that Spain is one of the European countries with the lowest labour participation of mothers of young children, as reported by Gutiérrez-Doménech (2005).

The remainder of the paper is organized as follows. In Section II we explain the details of the cash benefit policy under evaluation. In Section III we describe the data set and show empirical evidence for recent trends of female labour participation rates and, in particular, for those women potentially affected by the policy. In Section IV we formulate the econometric model and perform the DDD analysis. In Section V we show our estimation results. Finally, Section VI offers our conclusions.

II. The cash benefit for working mothers

According to Eurostat, public expenditure on policies related to family and children was about 0.5% of GDP in 2002 in Spain, in contrast to an average of 2.2% in the EU-15 countries.[1] In January 2003 several changes concerning family policy were undertaken in Spain (Law 46/2002). Some of these changes were aimed at promoting fertility and others at helping to reconcile family and work.

With the aim of promoting fertility, a reform of the income tax was implemented, with two main changes concerning the treatment of families with children: an increase in the tax deduction regarding the number of children and an increase in the tax deduction for each child aged under 3 years. The first of these changes was implemented as follows: before the policy, families could reduce their annual taxable income by 1200 euros for the first and second child and by 1800 for the third child and subsequent children. Since the policy was introduced, families have been able to reduce their annual taxable income by 1400 euros for the first child, 1500 euros for the second child, 2200 for the third child and, finally, 2300 for each subsequent child. The second change allows families with children aged under 3 years to increase the reduction of their taxable income from 300 to 1200 euros per child. Both of these innovations are aimed at promoting fertility.

With the aim of reconciling family and work, a monthly cash benefit for working mothers of children aged under 3 years was introduced. The monthly cash benefit amounts to 100 euros per child aged under 3 years. To be eligible, working mothers must fulfill certain conditions in relation to the number of hours worked. These conditions differ for full-time and part-time working mothers. In particular, full-time female workers must work at least 15 days per month. For part-time female workers the equivalent figure is 20 days. Furthermore, part-time female workers are eligible only if they work at least 50% of full-time hours. There are two alternative ways in which the benefit can be received. One possibility is that eligible females apply to the Public Administration to receive the cash benefit via bank transfer. The other possibility is

[1] See Instituto de Política Familiar (2006) for a comparison across European countries.

to deduce from their annual payable taxes an amount equal to the corresponding annual cash benefit. There is an upper limit to the cash benefit given by the annual social security payroll taxes. However, this policy is not subject to an income test to determine eligibility. The analysis of the policy effect on labor market participation of such mothers is the main objective of this paper.

How much child care can the cash benefit buy and how much does it represent in terms of females earnings? The cash benefit covers, on average, more than one third of the cost of private day-care centers in Spain, so it can finance a substantial amount of the child care costs faced by working mothers.[2] According to the average female earnings by education level reported by the Spanish Statistical Office in 2002, a cash benefit of 100 euros per month represents 13.0% of the earnings for females with elementary education, 7.5% if the woman has completed high school education and 5.2% in the case of college educated females. However, if instead of considering observed wages, we consider the potential wages of the whole female population these numbers would be higher given that the selection process brought about by the decision of women to join the labour market, is expected to leave out of the market those women with lower potential earnings.

In 2003, when this family policy was introduced, 625000 mothers benefited from it. In 2004, the figure was about 664000 mothers. According to the Institute of Family Policy, approximately six per cent of families in Spain benefit from the policy.

The simultaneous implementation of the cash benefit for working mothers of children aged under 3 years and the increase in the tax deduction for mothers of children aged under 3 years irrespective of their working status, prevents us from isolating the effect of each of the policy reform components on female labour market participation. As stated in the introductory Section, previous research for other countries shows that cash benefits for mothers, irrespective of their working status, can reduce female labour market participation by increasing intra-household specialization of work. So, in the case of the policy reform we are analyzing, the potential positive effect of the cash benefit on female labour market participation could be lessened by the other component of the policy reform, the increase in the tax deduction for all mothers of children aged under 3 years. However, in the case of the particular policy passed in Spain the tax deduction is not expected to have a strong effect on female labour market participation for two reasons. First, the tax deduction is not effective until one year after the earnings are made, which is especially important if borrowing restrictions against future disposable income operate in the credit markets. Second and more important, it implies a small increase in the annual disposable income for females.[3]

[2] In 2005, the average cost of private day-care centers in Spain was 254 euros per month (Consumer (2005)). Thus, a monthly cash benefit of 100 euros represents 39.37% of that cost.

[3] For example, for a female whose marginal income tax rate is 20%, the increase in the tax deduction from 300 to 1200 euros represents an increase in the annual disposable income of 180 euros, i.e, 15 euros per month.

III. Data and sample

The dataset we use in this paper is from the Labour Population Survey (EPA) conducted quarterly by the Spanish Statistical Office (INE). These data consist of pooled cross-sections of more than 150000 individuals, covering the period 1996-2004, i.e., seven years before and two years after the policy implementation.[4] We use a subsample of the raw data and, in particular, we focus on married females aged under 45.[5] Our focus is on this group of females, as single female labour participation rates are similar to male participation rates. Furthermore, the proportion of single mothers is still very low in Spain compared to other countries.[6] Left out of the analysis are those women belonging to the oldest cohorts, who will certainly have different labour behaviour and are not likely to be eligible for the policy.

As an initial approach to the question at hand we provide a descriptive analysis of the labour participation behaviour of our sample of Spanish women. Since there are substantial differences in participation across education groups, we distinguish three groups of individuals. The first group, which includes elementary educated females, represents 14.475% in our sample. The second group, composed of females having completed their high school education, is the most numerous one, representing 65.406% of the sample. The remaining 20.119% corresponds to the third group, those females who have completed college education.

Table 1, columns 2-5, provides the sample mean and standard deviation of several individual characteristics in 2004: the number of children, the fraction of women who worked in the previous year, the fraction of women whose spouse is working and woman's age. This information is reported by educational level for several groups of females: married females, married females who have no children, married mothers, married mothers of children aged 4 to 6 years old and finally, the targeted group of the policy, married mothers of children aged under 3 years. As expected, there is a negative relationship between the number of children and the educational level of the mother. As regards previous year employment status, the higher the educational attainment of the women, the higher the fraction of women who worked during the previous year. This fraction is very different across motherhood status, those exhibiting the lowest fraction being the mothers of young children, and especially so if they only have elementary education. With respect to the spouse's working status, no substantial differences are observed across motherhood

[4] In 2005 some changes were introduced in the Labour Population Survey. First, instead of providing the individual's age as the number of years or months since birth, Age is coded into intervals, being the first of them 0-4 years. This does not allow us to use data from 2005 onwards for our purposes since we cannot identify the targeted group of the policy, mothers of children aged under 3 years. Second, several changes were made concerning the questions aimed at determining the employment and participation status of the individual. As a result of this last change, the number of employed people increased and the number of unemployed people decreased in 2005 with respect to 2004.

[5] In fact, we understand as married females all females living in biparental households and being the head or the partner of the head of the household, irrespective of whether they are married or not.

[6] The average participation rate of married females aged under 45 is 65.52% for the period 2000-2004, whereas the average participation rate of single females is 89.93%. If we focus on mothers, the average participation rate of married females is 61.15%, whereas the average participation rate of single females is 86.51%. The average proportion of single mothers for the period 2000-2004 is 6.12%.

status. However, across educational groups, we observe that, the higher educational attainment is, the higher the fraction of women whose spouse is working, which reflects marital sorting by education. Finally, female's age is positively related with the age of the youngest child and mothers of children younger than 6 are older the higher their educational level is, which is a consequence of the differences in maternity ages across educational groups.

Table 1, column 6, shows the participation rates for 2004 across educational attainment and motherhood status. In the last two columns of Table 1 we show the annual average growth rate of participation for a period of six years before the policy implementation (1997-2002) and for the period of two years after the policy implementation (2003-2004). There are several remarks to be made. First, we observe a positive and strong correlation between education and participation and a negative correlation between motherhood and participation. More specifically, the lowest participation rates are found in the group of mothers of children aged under 3 years. However, differences in participation across motherhood status are more striking the lower female educational attainment is. Second, we observe an increase in the participation rates of married females throughout the period under consideration. This is part of the trend observed in Spain since the beginning of the eighties. Third, we observe that the increase in participation of married females is primarily due to the increase in the participation of mothers, as increases in non-mothers participation are minor, except for the elementary educated group in the period 1997-2002. Finally, if we focus on mothers of young chlidren, we find that mothers of children aged 4 to 6 years present a similar evolution before and after the policy implementation: the annual average participation growth rate is 2.38% in the before-policy period and 2.16% in the after-policy years. However, the targeted group of the policy, mothers of children aged under 3 years, shows a very different pattern: whereas the annual average participation growth rate is 1.61% in the before-policy period, this growth rate reaches 5.02% in the after-policy period. This empirical evidence points to a potential positive effect of the policy on the participation rate of the targeted group of females. Conditioning on educational attainment, the differences in the participation growth rate before and after the policy for the targeted group are more striking for those with elementary or high school education than for college educated mothers. As a final remark, for mothers of children aged 4 to 6 with elementary education, the annual average participation rate experiences an important increase in the after-policy period. This may be related to the fact that a fraction of these mothers, those whose youngest child is 4 years old in 2004, were policy-eligible in 2003. Then, their participation in 2004 could be indirectly affected by the policy through their decision to participate in the previous period, when they were policy-eligible.

In the next Section we formulate an econometric model in order to measure how much of the observed change in the participation rates of the eligible females can be attributed to the implementation of the policy.

TABLE 1: Descriptive statistics

	No. of children	% working last year	% working spouse	Age	Particip. rate (%)	Annual growth rate of avg. particip. (%) Before policy	After policy
I. Married females			2004			1997-2002	2003-2004
Elementary	1.92 (1.06)	34.00 (47.37)	85.65 (35.06)	38.40 (6.05)	49.75 (50.00)	2.27	2.53
High School	1.47 (0.91)	51.73 (49.97)	93.04 (25.43)	36.07 (5.93)	66.10 (47.34)	1.56	2.87
College	1.26 (1.00)	79.68 (40.24)	96.07 (19.44)	36.39 (5.39)	88.42 (32.01)	0.19	0.77
All	1.49 (0.97)	54.79 (49.77)	92.57 (26.22)	36.47 (5.90)	68.22 (46.56)	2.47	3.21
II. No mothers							
Elementary	-	46.34 (49.90)	81.92 (38.51)	34.75 (7.78)	62.99 (48.31)	2.69	0.85
High School	-	69.95 (45.85)	92.90 (25.68)	31.26 (6.30)	84.19 (36.49)	0.94	1.33
College	-	82.53 (37.98)	94.22 (23.34)	32.28 (5.01)	94.70 (22.40)	0.10	1.34
All	-	72.35 (44.73)	92.58 (26.21)	31.81 (6.11)	86.09 (34.61)	1.22	1.54
III. Mothers							
Elementary	2.09 (0.93)	32.89 (46.99)	85.98 (34.72)	38.73 (5.75)	48.57 (49.98)	2.12	2.42
High School	1.76 (0.69)	48.12 (49.97)	93.06 (25.41)	37.03 (5.36)	62.51 (48.41)	1.52	3.02
College	1.75 (0.73)	78.59 (41.02)	96.78 (17.66)	37.97 (4.65)	86.00 (34.70)	0.12	0.43
All	1.81 (0.75)	51.04 (49.99)	92.58 (26.22)	37.47 (5.35)	64.40 (47.88)	2.35	3.24
IV. Mothers of children between 4 and 6 years							
Elementary	2.30 (1.09)	30.00 (45.84)	85.97 (34.74)	35.06 (5.68)	48.66 (50.00)	-0.49	10.18
High School	1.76 (0.73)	44.72 (49.72)	93.61 (24.44)	34.94 (4.70)	61.31 (48.71)	1.68	1.06
College	1.81 (0.79)	77.94 (41.47)	97.07 (16.87)	37.42 (3.75)	86.04 (34.66)	-0.01	-0.14
All	1.83 (0.81)	49.66 (50.00)	93.43 (24.78)	35.45 (4.75)	64.79 (47.77)	2.38	2.16
V. Mothers of children under 3 years							
Elementary	2.16 (1.14)	22.07 (41.49)	86.40 (34.30)	31.89 (5.75)	38.46 (48.67)	0.59	7.35
High School	1.66 (0.74)	46.82 (49.90)	93.94 (23.87)	32.36 (4.78)	57.68 (49.41)	0.94	5.56
College	1.64 (0.74)	77.34 (41.86)	97.30 (16.20)	34.50 (3.86)	83.63 (37.01)	0.15	0.61
All	1.71 (0.80)	52.25 (49.95)	94.05 (23.66)	32.87 (4.78)	62.48 (48.42)	1.61	5.02

Note: As can be observed, in some cases the annual participation growth rate for all women without conditioning on the education level, is higher than the growth rate we observe for each educational group. This is not actually surprising if we notice that the weights corresponding to each education group change from one year to another. Standard deviations in parenthesis.

IV. Econometric specification

In this Section we perform a DDD estimation approach in order to measure the potential effect of the policy on the labour participation rate of those women affected by its implementation.

We estimate a probit model for the participation decision and we measure the potential effect of the policy based on the DDD estimator, broadly used in the literature of natural experiments or pseudo-experiments. This kind of analysis, which has a rich tradition in psychology, has been used in economics to study the effects of public policies on individual behaviour. The introduction of a policy change provides a natural experiment that allows the evaluation of the policy effect on an outcome of interest if we observe individual behaviour before and after its implementation. As there can be other factors, such as macroeconomic condition changes or any kind of trend, that influence the outcome of interest, the usual approach consists of examining this outcome not only for those groups affected by the policy change, but also for those similar groups that are not eligible for the policy but would presumably be subject to the other influences as well. In particular, in the case we are analysing two issues reinforce the importance of these considerations: (i) the cash benefit policy for working mothers is introduced simultaneously with certain increases in the tax deductions for children from the annual taxable income that benefits mothers in general; (ii) a decrease in the unemployment rate is observed in Spain over the last few years. This decrease could affect female participation in different directions. On the one hand, it can encourage them to search for a job. On the other, the lower probability of their husbands being unemployed can induce them to take a spell out of the labour market.

In the natural experiments literature, the policy-eligible group of the population is called the treatment group, and a suitable subset of the non-eligible group is used as a control group. The comparison of the change in the behaviour of the treatment group relative to the change experienced by the control group provides what is known as the 'difference-in-differences' (DD) estimator, which can be interpreted as an estimator of the policy effect. However, as pointed out by Meyer (1995), it may well be the case that the outcome of interest systematically evolves differently for the treatment and control groups and hence, the omission of a specific trend for the treatment group would bias the estimation of the policy effect. For this reason, we allow for different trends in the two groups and we build a DDD estimator by comparing changes in the behaviour of the treatment group with changes for the control group correcting for their potential different trends. This identification strategy allows us to capture the effects of the policy. Meyer (1995) provides an excellent survey of this kind of methodology, which has been applied in a huge variety of contexts[7].

In this paper, the public policy to be evaluated is the introduction of a cash benefit for working mothers of children aged under 3 years. The treatment group is composed by mothers of children aged under 3 years. In our analysis we consider as the control group mothers of children aged under 6 years that are non-eligible for the policy, i.e., mothers of children aged 4 to 6 years. We restrict the control group to include only mothers of children aged under 6 as they

[7] See for example Card and Krueger (1994), Meyer, Viscusi and Durbin (1995), Cummins, Hassett and Hubbard (1994) and more recent works as Stephens and Ward-Batts (2004).

are expected to have a closer behaviour to the treatment group. The time and care demands of children younger than 6 are very different from those of older children, given that compulsory education in Spain starts at 6 years. Furthermore, we eliminate from the control group those non-policy eligible mothers of children aged under 6 years who benefited from the policy in a previous period. In particular, these are mothers whose youngest child was 4 years-old in 2004. The inclusion of these women in the control group could underestimate the direct effect of the policy on the participation of mothers of children aged under 3. This will be the case if there is an indirect effect of the policy such that the potentially higher participation of policy-eligible mothers translates into a higher participation once their children get over the age of 3 and they become non-policy-eligible. In fact, this could be an important positive effect of the policy, but which we do not measure here.[8]

Let y_{it} be a binary variable which takes on the value 1 for those females who participate in the labour market and 0 otherwise.

$$\Pr(y_{it} = 1 | After_t, Treatment_i, x_{it}) =$$
$$\Phi\left(\alpha + \beta_1 After_t + \beta_2 Treatment_i + \beta_3 After_t * Treatment_i + \beta_4 t + \beta_5 t * Treatment_i + x'_{it}\gamma\right) \tag{1}$$

for $i = 1, ..., N$ and $t = 1, ..., T$, where i indexes females, t indexes time and $\Phi(.)$ stands for the cumulative normal distribution function. The variable $After_t$ is a dummy variable which equals 1 after the introduction of the policy, $Treatment_i$ is a dummy variable which equals 1 for mothers of children aged under 3 years, i.e., the treatment group, and thus, the interaction $After_t * Treatment_i$ can be interpreted as the 'policy variable'. The vector x_{it} contains explanatory variables related to socioeconomic and family characteristics. As regards the time trend, two issues are worth noting. First, we allow a different linear time trend for the treatment and control groups so that we do not ascribe to the policy an effect that is in fact related to systematic different behaviour between the treatment and the control group along time. Second, once this issue has been taken into account, our identification assumption is that there are no other contemporaneous shocks, except for the policy under analysis, that could affect the relative behaviour of the treatment and control groups over time.[9]

The variables $After_t$ and $Treatment_i$, as well as the interaction variable, $After_t * Treatment_i$, have a useful interpretation in terms of labour participation for the treatment and control groups before and after the policy. The variable $After_t$ controls for differences in labour participation before and after the implementation of the policy for both groups. There are many possible changes in the socioeconomic environment that can affect female behaviour, such as changes in other dimensions of the family policy scheme, in the returns to labour market experience or in the uncertainty affecting husband's earnings or employment[10]. In particular, in this case the

[8] Furthermore, we recognize that even those females that are not mothers of children younger than 3 years could potentially be affected by the policy as they can decide to have a newborn. However, we think it is too early to find evidence of such effects. Thus, we have concentrated on the effect of the policy on participation decisions and postpone the analysis of its effect on fertiliy decisions for future work.

[9] Our specification in this concern is similar to the one in Francesconi and Van der Klaauw (2007).

[10] See for example Olivetti (2003) and Attanasio et al. (2004, 2005) for the importance of these issues on the

simultaneous introduction of the cash benefit for working mothers and the increase in the tax deductions for children from the annual taxable income justifies the inclusion of this variable in the model specification. The variable $Treatment_i$ controls for systematic differences in the participation behaviour of the treatment and control groups. The interaction variable $After_t * Treatment_i$ controls for the relative change in labour participation in the treatment group after the policy. We include a time trend common to both groups and a specific trend for the treatment group. The justification for considering a specific trend for mothers of young children is based on evidence for other countries like the US where recent changes in female labour participation are concentrated among this group of mothers, see for example Attanasio et al. (2004).

Among the explanatory variables in x_{it} we consider variables related to individual and household characteristics. The educational attainment is controlled by the binary variables $College$ $education$, which equals 1 for females with college education and $High\ School\ education$, which equals 1 for females with high school education completed. Elementary educated females are the reference category. The age of the female in years is measured by the variable Age. We also include a quadratic term of this variable in order to allow for nonlinear effects. The presence of children in the household is represented by the variable $Number\ of\ Children$. Female employment status in the previous year is controlled by a binary variable, $Work_{-1}$, which equals 1 if the female worked.[11] As is argued by De la Rica and Ferrero (2003) and others, past employment is likely to affect current participation positively through different channels; for example, females who have already been in the labour market do not have to incur the costs derived from entering it (such as the costs related to searching for a job). Finally, the employment status of the spouse is controlled by the binary variable $Spouse\prime s\ work$, which equals 1 if he is working.

One important methodological issue is that the potential endogeneity of fertility in the participation equation could produce biased results. However, the main problem that arises in this context is the difficulty of finding accurate instruments for fertility in the participation equation, in the sense that they should be correlated with fertility but not correlated with unobservable factors affecting participation. Some of the works which have adopted this strategy have constructed instruments based on the existence of twins in the first birth or on the sibling sex composition in families with two or more children.[12] In general, the use of these kind of instruments restrict the sample to women who have already had children. As is stressed in some of these works, it is not clear that the results can be extended to a general context of the impact of childbearing on participation. An additional drawback of some of these papers is to treat fertility and participation variables as continuous. A suitable approach for dealing with discrete variables is Manski et al. (1992). They propose a bivariate model in which they jointly estimate the equation of interest and the equation for the variable for which the endogeneity is being treated. This

determination of female participation.

[11]The inclusion of the variable $Work_{-1}$ could be enough to control for the indirect effect of the policy on participation of mothers of children aged over 3. However, this variable is an imperfect way of controlling the effect of previous labour market experience on the participation decision (total labour market experience is not available in the survey). For this reason, we choose to exclude from the control group those who are mothers of 4 years old children in 2004, who benefited from the policy in 2003.

[12]See Rosenzweig and Wolpin (1980), Bronars and Grogger (1994), Gangadharan and Roseenbloom (1996) and Angrist and Evans (1998).

strategy has been used in the context of female labour participation and fertility decisions by De la Rica and Ferrero (2003) and Carrasco (2001). Furthermore, studies which have focused on the effect of fertility on participation decisions have reported results of opposite sign.[13] Given these difficulties, a common approach in several works in the literature that analyze issues related with female participation decisions is to assume exogenous fertility.[14] We follow the same strategy here.

Finally, there is a last methodological issue we should mention. We use pooled cross sections, as panel data are not available for the period of interest. Then, we cannot control for unobserved heterogeneity through the use of person-specifc fixed effects.

Construction of the DDD estimator

The nonlinearity of the model, which comes from the outcome of interest being a binary variable, makes the construction of the DDD estimator slightly different from the linear case. The effect of the policy is not given by the estimate of a single parameter, as it would be in a linear model, but by differences in estimated probabilities. Let $\Delta p1\_aft$ be the change in the participation probability experienced by the treatment group $k-1$ periods after the policy with respect to this probability one year before the policy. Let $\Delta p0\_aft$ be the analogous change in the participation probability for the control group. Let $\Delta p1\_bef$ be the change in the participation probability experienced by the treatment group one period before the policy with respect to this probability $k+1$ periods before the policy. Let $\Delta p0\_bef$ be the analogous change in the participation probability for the control group. These differences can be written as:[15]

$$\begin{aligned} \Delta p1\_aft &= \Pr\left(y_{i,k^*-1+k}=1 \mid Treatment_i=1\right) - \Pr\left(y_{i,k^*-1}=1 \mid Treatment_i=1\right) \\ \Delta p0\_aft &= \Pr\left(y_{i,k^*-1+k}=1 \mid Treatment_i=0\right) - \Pr\left(y_{i,k^*-1}=1 \mid Treatment_i=0\right) \\ \\ \Delta p1\_bef &= \Pr\left(y_{i,k^*-1}=1 \mid Treatment_i=1\right) - \Pr\left(y_{i,k^*-1-k}=1 \mid Treatment_i=1\right) \\ \Delta p0\_bef &= \Pr\left(y_{i,k^*-1}=1 \mid Treatment_i=0\right) - \Pr\left(y_{i,k^*-1-k}=1 \mid Treatment_i=0\right) \end{aligned} \tag{2}$$

where k^* is the inception date of the policy ($k^*=2003$ in our case). The available data allows us to perform the analysis for $k=1$, 2. From these differences of probabilities, we build the DDD estimator as follows:

$$DDD = (\Delta p1\_aft - \Delta p1\_bef) - (\Delta p0\_aft - \Delta p0\_bef) \tag{3}$$

We perform the analysis considering $k=2$, which implies a comparison of changes in participation probabilities from 2002 to 2004 with respect to changes from 2000 to 2002 for both the treatment and the control group. Thus, we are comparing for both groups the evolution of the participation probabilities in periods of the same length.

[13] Carrasco (2001) and De la Rica and Ferrero (2003) find that the estimated effect of fertility on participation is downward biased when fertility is assumed to be exogenous, whereas Angrist and Evans (1998) and Angrist (2001) find an upward bias.

[14] For example, Wellington (2001) investigates self-employment as a strategy to balance family and career and Gutiérrez-Doménech (2005) studies women´s transitions from employment to nonemployment after first birth in several countries of the EU. They both assume exogenous fertility.

[15] In order to simplify notation, we omit $After_t$, x_{it} and the linear trend from the vector of conditioning variables.

V. Estimation results

In this Section we present the results for the model specification that was explained above in order to isolate the effect of the cash benefit policy introduced in Spain in 2003. We consider data covering the years 1996 to 2004, i.e., seven years before and two years after the implementation of the policy.

The estimation results of equation (1) are shown in Table 2. All the explanatory variables we include are significant at 5%, except for variable $Trend$ and Age related variables, which are non-significant at the usual levels. As regards the variables related to the structural behaviour of females, the sign of the coefficients are supported by microeconomic theory.[16] First of all, the variable $Number\ of\ Children$ has a negative effect on participation as was expected. Furthermore, the variable controlling for children aged under 3 years, $Treatment$, has a negative sign coefficient, reflecting the negative effect of the presence of young children on woman's participation. Concerning the educational level, since elementary education is the reference category, the sign of the coefficients of the high school and college education variables are the expected ones. That is to say, the higher the education the higher the participation probability. This fact reflects that the opportunity cost of being at home is an increasing function of educational level. The variable controlling for previous year employment status, $Work_{-1}$, has a positive effect on participation, along the lines of De la Rica and Ferrero (2003). The variable related to the employment status of the spouse, $Spouse's\ work$, has a negative coefficient that can be related to the so-called 'added worker effect': females married to unemployed males are more likely to participate in the labour market. The negative coefficient of the variable $Trend * Treatment$ reflects that the trend is flatter for the treatment group. The dummy variable $After$ has a positive effect on female labour participation. This reflects the relatively higher participation rates of all women in the years after the policy implementation. Finally, we focus on the variable that provides us with some insights about the question we pose in this work. The 'policy variable', $After * Treatment$, has a positive and significant effect on the participation decision. This means that there is a particular behaviour of the targeted group, mothers of children aged under 3 years, that manifests itself via a relatively higher increase in participation with respect to the increase for the control group after the policy implementation.

[16]See, for example, Becker (1965) and Browning (1992).

TABLE 2: Participation equation probit estimates

After	0.057 (0.0179)
Treatment	-0.186 (0.0244)
After*Treatment	0.071 (0.0245)
Trend	-0.004 (0.003)
Trend*Treatment	-0.012 (0.0048)
Number of children	-0.112 (0.0048)
Work_{-1}	1.883 (0.0083)
High School education	0.138 (0.0103)
College education	0.662 (0.0136)
Spouse's work	-0.371 (0.0145)
Age	0.007 (0.0077)
Age^2	-0.0001 (0.0001)
Constant	-0.089 (0.1291)
Log-likelihood	-76902.65
Pseudo-R^2	0.3511
Number of observations	174155
In the treatment group:	97696
After the policy:	27385
Before the policy:	70311
In the control group:	76459
After the policy:	18850
Before the policy:	57609

Note: Robust-heteroskedasticity standard errors in parenthesis.
Treatment group: mothers of children aged under 3 years.
Control group: mothers of children aged 4 to 6 years, excluding those whose youngest child was 4 years old in 2004.

From our probit estimates, we compute the estimated participation probabilities for the treatment and control groups and we obtain the DDD estimate. As we mentioned above, in order to calculate the DDD we compare for both groups the change in the participation probabilities from 2002 to 2004 with the change from 2000 to 2002, as was stated in (2) and (3). The data from 1996 to 1999 that are also included in our sample only serve the purpose of estimating the time trend, but we compare the treatment and control groups in periods of the same length. The results can be seen in Table 3. The variation in the estimated participation probability for the control group before the policy implementation, $\Delta p0\_bef$, is around -0.272 percentage points, while this variation after the policy implementation, $\Delta p0\_aft$, is around 1.719 percentage points. For the treatment group it is remarkable that, whereas the variation before the policy implementation, $\Delta p1\_bef$, is -1.226 percentage points, the variation after the policy is implemented, $\Delta p1\_aft$, is around 3.690 percentage points. These results give an estimate of the DDD of 2.926 percentage points, an increase which could be potentially ascribed to the family policy. This variation represents 5.164% of the average labour market participation of the treatment group in 2002, the year before the policy implementation.

We have also performed the analysis without excluding from the control group those women in the sample whose youngest child was 4 years old in 2004 (i.e, women that were policy-eligible in 2003). We find that the estimated DDD is slightly lower with this extended sample, 5.033%.[17] According with this result, the indirect effect of the policy on the group of women that are not currently policy-eligible could be partially controlled by the variable $Work_{-1}$. However, we think it would be necessary to have more after-policy observations to further investigate this indirect effect of the policy.

TABLE 3: Estimated changes in participation probabilities and DDD estimate

	Δp_aft	Δp_bef	Difference	DDD estimate	% of increase over treatment group participation rate in 2002
Treatment group ($\Delta p1$)	3.690 (0.490)	−1.226 (0.254)	4.916 (0.651)	2.926 (0.906)	5.164%
Control group ($\Delta p0$)	1.719 (0.484)	−0.272 (0.253)	1.991 (0.632)		

Note: Standard errors in parentheses computed through the Delta method

The descriptive analysis in Section III shows important differences in participation probabilities across educational levels. We have obtained DDD estimates by educational attainment. The results are shown in Table 4. As can be seen, the potential effect of the policy is more striking for those with elementary or high school education than for those with college education. In particular, the DDD estimate represents 8.690% of the average labour market participation in 2002 for those females in the treatment group with elementary education, 5.670% for those with high school education and 3.079% for those with college education. This is reasonable as females with elementary or high school education were not only less likely to participate in the labour market before the policy implementation than college educated females, but also the cash benefit represents a higher proportion of their potential earnings. Therefore the benefit can be more relevant in determining the participation of females with lower education.

TABLE 4: DDD estimates by education level

	DDD estimate	% of increase over treatment group participation rate in 2002
Elementary	2.902 (0.959)	8.690%
High School	2.935 (0.931)	5.670%
College	2.544 (0.704)	3.079%

Note: Standard errors in parentheses computed through the Delta method.

[17]Estimation results are available upon request.

Our estimation results allow us to say something about the elasticity of labour market participation of the treatment group with respect to child care prices. If we assume that there is no effect of the cash-benefit on the equilibrium price of child care services, we can say that the elasticity of labour market participation of the treatment group with respect to child care prices implied by our DDD estimate is -0.131. This number is in the range estimated by Anderson and Levine (1999).[18]

As stated above, the potential endogeneity of the fertility decision could affect the estimation results. In spite of the difficulties in finding appropriate instruments for the fertility decisions that was explained in Section IV, we follow the approach suggested by Manski et al. (1992) to estimate a bivariate probit model for the participation and fertility decisions.[19] We find that the potential positive effect of the policy is robust to this alternative specification, altough it is underestimated when assuming exogenous fertility. In any case, this result should be taken with caution given the difficulties we mention above.[20]

VI. Conclusions

The objective of this work is to analyze the effects of a public policy introduced in Spain in 2003 aimed at reconciling family and work. The policy provides working mothers of children aged under 3 years with a monthly cash benefit of 100 euros per child. Other studies have investigated the effect of similar policies in Norway (see Schone (2004) and Naz (2004)) and United Kingdom (see Francesconi and Van der Klaauw (2007)). Our study is a contribution to that literature, with the additional interest of it concerning one of the countries with the lowest labour participation of mothers, as reported by Gutiérrez-Doménech (2005).

The methodological approach we use in this work is based on the recent natural experiments literature, which has been broadly used in the evaluation of public policies in a wide variety of contexts. We formulate an econometric model of labour participation and measure the potential effect of the policy by means of a DDD estimator.

Our estimation results point to a small positive effect of the policy in increasing labour participation of the targeted group, mothers of children aged under 3 years. After the policy was implemented, this group of mothers increases its labour market participation 2.926 percentage points more than the non-policy-eligible females. This estimated relatively higher increase rep-

[18] They found that for mothers of children aged under 13, the elasticity of labor market participation with respect to the market price of child care services ranges between -0.05 and -0.35. They also found this elasticity being relatively higher for low educated mothers as well as for mothers of children aged under 6.

[19] Variables related to the innate ability of females to reproduce, such as the number of brothers and sisters of the female (as a proxy for her genetic ability to reproduce), or variables related to religion could be used as instruments. However, these variables are not usually reported in the Labour Force Surveys. In the bivariate probit model estimated by us we have considered the number of children in the household apart from the newborn and the interaction of education and age variables.

[20] Estimation results of the bivariate model are available upon request.

resents 5.164% of the participation rate the targeted group exhibited in 2002, the year before the policy was implemented. It should be noted, though, that this effect is not homogeneous across educational groups, it being stronger for elementary and high school educated females. More specifically, for elementary educated females the policy estimated effect represents 8.690% of the participation rate in 2002. The same effect for high school educated females is 5.670%. Finally, for college educated females, the effect is less remarkable, it being only 3.079% of the participation rate in 2002.

We conclude from our analysis that the observed variation of participation of mothers of children aged under 3 years could be partially ascribed to the cash-benefit policy that came into force in 2003. Our results have implications for governments willing to design family policies aimed at increasing labour market participation of mothers of young children. As a future research project it will be interesting to measure the indirect effect of the policy on the participation of mothers of children aged over 3 resulting from their stronger attachment to the labour market during the early maternity ages. To perform such an analysis it will be necessary to observe policy-eligible women for several years after the period in which they benefit from the policy.

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