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 risen by 3 percentage points compared to the rate for non-policy-eligible females. This represents 5% of their average labour market participation in 2002, the year before the policy was implemented. This overall

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policy effect is dominated by the effect of the policy among high school educated females.

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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 mothers' labour force participation 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 Sweden, 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 there is reason to believe that women experience 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 employment among women following the first birth. Secondly, 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). Thirdly, 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%, in sharp contrast to the average 25% for the European countries. Fourthly, the Spanish Labour Population Survey (2004) clearly indicates that family responsabilities are an important reason why Spanish females do not participate in the labour market. For the group of people aged under 45 years, less than 1% of males report family responsibilities as the main reason for not participating in the labour market; however, this figure is 29% in the case of females. 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 gender gap labour force participation. As a group, mothers of young children have among the lowest labour participation rates, see for example OECD

(2002). In addition, 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 are an 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 objective of this paper is to measure the direct effect of this policy on the participation of mothers of children aged under 3 years, and thus, to contribute to the discussion about whether to increase the amount of public resources devoted to promoting 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 an indirect effect on 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.

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 rate of change 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 rate of change in 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 find 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 3 percentage points higher than the increase experienced by non-eligible females. This represents 5% of the before-policy average participation rate of the policy-eligible women. Concerning the potential heterogeneity of the effect across education levels, we find a positive and significant effect of the policy for high school educated females, that we do not find for females with elementary or college education. Thus, the overall effect of the policy is dominated by the effect stemming from the high school educated group.

Our paper is related to others in the literature analyzing family policies. 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. 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 below 1% of GDP in 2002 in Spain, in contrast to an average of 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 each 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 sub-

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

sequent child. The second change raises the additional tax deduction for having a child under 3 years 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 to deduct 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, there is no an income test to determine eligibility for the benefit. The analysis of the policy effect on labor market participation of mothers of children aged under 3 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% of the earnings for females with elementary education, 8% if the woman has completed high school education and 5% 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 process of selection into the labour market is expected to exclude those women with lower potential earnings.

In 2003, when this family policy was introduced, there were 625000 working mothers who collected the cash benefit; as mentioned above, all working mothers with children under 3 years were eligible. 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 the separate policy reform components on female labour market participa-

[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% of that cost.

tion. 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 available to all mothers 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 there are restrictions on borrowing against future disposable income in the credit markets. Second and more important, the tax deduction implies a small increase in the annual disposable income for females.[3]

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 restrict our attention to 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.

[3] Marginal income tax rates in Spain ranged between 15% and 45% in 2003 and the average marginal tax was around 20%. For a female whose marginal income tax rate was 20%, the increase in the tax deduction from 300 to 1200 euros represented an increase in annual disposable income of 180 euros, i.e, 15 euros per month.

[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, the first of them being 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 in 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 66% for the period 2000-2004, whereas the average participation rate of single females is 90%. If we focus on mothers, the average participation rate of married females is 61%, whereas the average participation rate of single females is 87%. The average proportion of single mothers for the period 2000-2004 is 6%.

As an initial approach to the question at hand we provide a descriptive analysis of the labour participation behaviour of our sample of married 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 15% of our sample. The second group, composed of females having completed their high school education, is the most numerous one, representing 65% of the sample. The remaining 20% corresponds to the third group, those females who have completed college education.

Table 1 provides the sample mean and standard deviations for continuous variables for 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, woman's age and the participation rate. Finally, the last column of Table 1 includes the sample sizes. 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 3 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 status. However, across educational groups, we observe that, the higher educational attainment, the higher the fraction of women whose spouse is working, reflecting marital sorting by education. 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, a consequence of the differences in maternity ages across educational groups. Finally, we observe a strong and positive 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.

TABLE 1: *Summary statistics*

	No. of children	% Working last year	% Working spouse	Age	Particip. rate (%)	Sample size
I. Married females			2004			
Elementary	1.92 (1.06)	34.00	85.65	38.40 (6.05)	49.75	10173
High School	1.47 (0.91)	51.73	93.04	36.07 (5.93)	66.10	45966
College	1.26 (1.00)	79.68	96.07	36.39 (5.39)	88.42	14139
All	1.49 (0.97)	54.79	92.57	36.47 (5.90)	68.22	70278
II. Not mothers						
Elementary	-	46.34	81.92	34.75 (7.78)	62.99	835
High School	-	69.95	92.90	31.26 (6.30)	84.19	7607
College	-	82.53	94.22	32.28 (5.01)	94.70	3926
All	-	72.35	92.58	31.81 (6.11)	86.09	12368
III. Mothers						
Elementary	2.09 (0.93)	32.89	85.98	38.73 (5.75)	48.57	9338
High School	1.76 (0.69)	48.12	93.06	37.03 (5.36)	62.51	38359
College	1.75 (0.73)	78.59	96.78	37.97 (4.65)	86.00	10213
All	1.81 (0.75)	51.04	92.58	37.47 (5.35)	64.40	57910
IV. Mothers of children aged 3 to 6 years						
Elementary	2.30 (1.09)	30.00	85.97	35.06 (5.68)	48.66	1340
High School	1.76 (0.73)	44.72	93.61	34.94 (4.70)	61.31	7960
College	1.81 (0.79)	77.94	97.07	37.42 (3.75)	86.04	2321
All	1.83 (0.81)	49.66	93.43	35.45 (4.75)	64.79	11621
V. Mothers of children under 3 years						
Elementary	2.16 (1.14)	22.07	86.40	31.89 (5.75)	38.46	1404
High School	1.66 (0.74)	46.82	93.94	32.36 (4.78)	57.68	8824
College	1.64 (0.74)	77.34	97.30	34.50 (3.86)	83.63	3598
All	1.71 (0.80)	52.25	94.05	32.87 (4.78)	62.48	13826

Note: Authors' calculations using data from the Labour Population Survey (EPA) conducted by the Spanish Statistical Office (INE). Sample married females under age 45. Standard deviations for continuous variables in parentheses.

Table 2 shows the average annual growth rate of participation for a period of six years before the policy implementation (1997-2002) and for a period of two years after the policy implementation (2003-2004). 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. 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. Focusing on mothers, those who have children aged 3 to 6 years present a similar evolution before and after the policy implementation: the annual average participation growth rate is around 2% in the before-policy and after-policy periods. However, the group targeted by the policy, mothers of children aged under 3 years, shows a very different pattern: whereas the average annual participation growth rate is around 2% in the before-policy period, this growth rate reaches 5% in the after-policy period. Conditioning on educational attainment, the differences are more striking for those with elementary or high school education than for college educated mothers. In order to know if the difference between the average participation growth rate in the before and after policy periods is significant, we run a regression of the annual participation growth rate on a dummy variable controlling for the after policy period. We find the difference to be significant at 5% in those cases that we mark with an asterisk in Table 2. In particular, this is the case for mothers of children aged 3 to 6 years with elementary education and for mothers of children under 3 years, both for the whole group and for those with high school education[7]. This empirical evidence points to a potential positive effect of the policy on the participation rate of the targeted group of females. As a final remark, for mothers of children aged 3 to 6 years with elementary education, the average annual 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 3 years old in 2004, were policy-eligible in 2003. Thus, 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.

[7] This test should be taken with caution as we only have 8 observations to perform it. If panel data were available, a different approach could be adopted as we could have used individual information. Unfortunately, this is not the case, and the test can only exploit aggregate information.

TABLE 2: *Variation in Participation Across Years*

	Average annual growth rate (%)	
	Before policy	After policy
I. Married females	1997-2002	2003-2004
Elementary	2.27	2.53
High school	1.56	2.87
College	0.19	0.77
All	2.47	3.21
II. Not mothers		
Elementary	2.69	0.85
High school	0.94	1.33
College	0.10	1.34
All	1.22	1.54
III. Mothers		
Elementary	2.12	2.42
High school	1.52	3.02
College	0.12	0.43
All	2.35	3.24
IV. Mothers of children aged 3 to 6 years		
Elementary	−0.49	10.18*
High school	1.68	1.06
College	−0.01	−0.14
All	2.38	2.16
V. Mothers of children under 3 years		
Elementary	0.59	7.35
High school	0.94	5.56*
College	0.15	0.61
All	1.61	5.02*

Note: Authors' calculations. The annual participation growth rate for all women may be higher than those observed for each educational group because the weights corresponding to each education group change from one year to another. We mark with an asterisk those cases in which the annual participation growth rate after the policy implementation is significantly different (at 5%) than in the before-policy period.

IV. Econometric specification

In this Section we formulate an econometric model in order to measure the effect of the change in policy on the labour force participation rate of those women affected by its implementation.

We estimate a probit model of 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 analyzing two issues reinforce the importance of these considerations: (i) the cash benefit policy for working mothers is introduced simultaneously with increases in income tax deductions for children that benefit mothers in general; (ii) a decrease in the unemployment rate is observed in Spain over the period. The decrease in unemployment could affect participation of females in different directions. On the one hand, it could encourage them to search for a job. On the other, the lower probability of their husbands being unemployed could 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 different underlying 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 large variety of contexts[8].

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

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, who constitute the treatment group. 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 3 to 6 years. We restrict the control group to include only mothers of children aged under 6 as they 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 3 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 one that we do not measure here.[9]

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

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

for $i = 1, ..., N$ and $t = 1, ..., T$, where i indexes individuals, 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. 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 differences in the evolution of the behaviour between the treatment and the control group over time. 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.[10]

[9]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.

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

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 husband's earnings or employment[11]. In particular, in this case the simultaneous introduction of the cash benefit for working mothers and the increase in the income tax deductions for children 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. As noted above, 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}, 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.[12] 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 $Spouse's\ work$, which equals 1 if he is working.

Our specification treats fertility as exogenous with respect to participation. However, it could be the case that decisions about fertility and participation are made simultaneously, suggesting the use of instruments to account for the endogeneity of

[11] See for example Olivetti (2003) and Attanasio et al. (2004, 2005) for the importance of these issues for the determination of female participation.

[12] 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 3 years old children in 2004, who benefited from the policy in 2003.

fertility. The main problem that arises in this context is the difficulty of finding appropriate instruments for the fertility in the participation equation, i.e. instruments that are correlated with fertility but not correlated with unobservable factors affecting participation. Studies that treat fertility as endogenous do not yield consistent results. For example, 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. Given these difficulties, we have followed other papers in the literature that analyze female participation decisions assuming exogenous fertility.[13] However, as reported later, our results do not appear to be sensitive to this choice.

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 individual-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. The available data allows us to perform the analysis for $k = 1$, 2. We build the DDD estimator as follows:

$$DDD = (\Delta p1\_aft - \Delta p1\_bef) - (\Delta p0\_aft - \Delta p0\_bef) \tag{2}$$

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

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 3. All the explanatory variables we include are significant at the 5% level, except for variable $Trend$ and the Age related variables, which are non-significant at the usual levels. As regards the variables related to the structural behaviour of females, the signs of the coefficients are supported by microeconomic theory.[14] First of all, the variable $Number\ of\ Children$ has a negative effect on participation as 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 reflects the fact 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', that is, 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 in participation 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.

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

TABLE 3: *Participation equation probit estimates*		
After	0.06	(0.02)
Treatment	−0.19	(0.02)
After*Treatment	0.07	(0.02)
Trend	−0.04	(0.003)
Trend*Treatment	−0.01	(0.005)
Number of children	−0.11	(0.005)
Work_{-1}	1.88	(0.008)
High school education	0.14	(0.01)
College education	0.66	(0.01)
Spouse's work	−0.37	(0.01)
Age	0.07	(0.008)
Age^2	−0.01	(0.0001)
Constant	−0.09	(0.13)
Log-likelihood	−76902.65	
Pseudo-R^2	0.35	
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 parentheses.
Treatment group: mothers of children aged under 3 years.
Control group: mothers of children aged 3 to 6 years, excluding those whose youngest child was 3 years old in 2004.
Variables Trend, Age and Age^2 are divided by 10 in the estimation.

From our probit estimates, we compute the estimated participation probabilities for the treatment and control groups and we obtain the DDD estimate. As 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. The data from 1996 to 1999 that are also included in our sample only serve the purpose of estimating the time trend. The results can be seen in Table 4. The change in the estimated participation probability for the control group in the two years before the policy implementation, $\Delta p0\_bef$, is below 1 percentage point, while this change in the period 2002-2004, $\Delta p0\_aft$, is around 2 percentage points. For the treatment group, whereas participation falls in the period 2000-2002, ($\Delta p1\_bef$, represents a reduction of 1 percentage point), the change in the period 2002-2004, $\Delta p1\_aft$, is around 4 percentage points. These results give an estimate of the DDD of around 3 percentage points, an increase which we ascribed to the family policy. This variation represents 5% of the average labour market participation of the treatment group in 2002, the year before the policy implementation. Thus, since the cash benefit represents 39% of the average child care cost, the implied elasticity of labour market participation of the treatment group with respect to child care prices is -0.07, assuming that there is no effect of the cash-benefit on the equilibrium price of child care services. This number is in the range estimated by Anderson and Levine (1999).[15]

It is interesting to compare the estimated effect of the policy in Spain with the effect of similar policies in other countries. Francesconi and van der Klaaw (2007) found that an increase in the generousity of "in work" benefits for low income families with children in the United Kingdom in 1999 led to an increase in the single mothers' employment rate of about 5 percentage points. Although the magnitude of the credit increase was not homogeneous across mothers,[16] the reform implied an average benefit increase of 8 pounds per week, representing around 13% of the weekly child care cost, generating a response substantially higher than in the policy we analyze in this paper. We can think of two possible explanations. First, according to several studies,[17] a higher effect of this kind of policies is expected among single mothers than among married mothers. Second, the targeted group in Francesconi and van der Klaaw (2007) is low income families only, for which participation is more likely to be responsive.[18]

TABLE 4: *Estimated changes in participation probabilities and DDD estimate*

	Δp_aft	Δp_bef	Difference	DDD estimate	% Increase over treatment group participation rate in 2002
Treatment group $(\Delta p1)$	3.69 (0.49)	-1.23 (0.25)	4.92 (0.65)	2.93 (0.91)	5.17%
Control group $(\Delta p0)$	1.72 (0.48)	-0.27 (0.25)	1.99 (0.63)		

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 so we want to explore the potential heterogenous effect of the policy across education groups. We show DDD estimates

[15] 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.

[16] The reason is that it affected several parameters of the program at the same time: the amount of weekly credit to which the family was entitled depending on the age of the children and the income threshold to be entitled to the policy, among other issues.

[17] See, for example, Blau and Robbins (1988), Connelly (1992) or Connelly and Kimmel (2000).

[18] Unfortunately, we can not undertake the analysis conditioning on household income as the Spanish Labour Force Survey does not include such information.

by educational attainment in Table 5.[19] We cannot reject the null hypothesis of no policy effect for elementary and college education groups. For high school educated females we find a positive and significant effect of the policy. The DDD estimate is 3.06 percentage points, which represents an increase of the labor participation of the treatment group in 2002 of around 6%. The overall effect that we report in Table 4 is dominated by the effect of the policy for high school educated females.

TABLE 5: *DDD estimates by education level*

	DDD estimate	% Increase over treatment group participation rate in 2002
Elementary	1.74 (2.33)	5.22%
High school	3.06 (1.13)	5.91%
College	1.68 (1.05)	2.03%

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

As a check on the sensitivity of our findings, we have also performed the analysis without excluding from the control group those women in the sample whose youngest child was 3 years old in 2004 (i.e, women that were policy-eligible in 2003). We find that the estimated DDD is very similar with this extended sample, around 5% of the average labour market participation of the treatment group in 2002, the year before the policy implementation.[20] 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.

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, as a sensitivity check, we follow the approach suggested by Manski et al. (1992) to estimate a bivariate probit model for the participation and fertility decisions.[21] They propose a bivariate model in which the equation of interest and the equation for the variable for which the endogeneity is being treated are estimated jointly and apply this strategy to analyze the probability of high school graduation as a function of the family structure. In our bivariate probit model, we have considered the number of children in the

[19] These numbers are based on separate labor force participation equations estimated for the different education groups.

[20] Estimation results are available upon request.

[21] 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.

household apart from the newborn and the interaction of education and age variables as instruments. We find that the potential positive effect of the policy is robust to this alternative specification and indeed a bit larger than if exogenous fertility is assumed. In any case, this result should be taken with caution given the difficulties we mention above.[22]

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 3 percentage points more than the non-policy-eligible females. The largest point estimate and the largest implied proportional effect is that for women with a high school education, the only group for which the estimated effect is statistically significant.

We conclude from our analysis that the observed growth in 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 seeking 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.

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

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