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Childcare costs and Spanish mothers' labour force participation*

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Abstract

In Spain, female labour force participation is among the lowest in Europe. This paper analyzes the extent to which female labour force participation is affected by the cost of formal childcare. Both decisions, labour force participation and formal childcare use, are jointly considered by means of a bi-variate probit model that accounts for the sample selection. Based on data from the Spanish Time Use Survey, the study indicates that Spanish mothers' labour force participation is very elastic to changes in childcare costs.

Keywords: Childcare costs, female labour participation.

JEL classification: J13, J22, C35

1. Introduction

As in other Southern European countries, labour force participation of Spanish women has increased greatly during the last two decades. Participation rates have risen from about 42% in 1990 to more than 61% in 2006 (OECD, 2007a). Part of this variation is due to the considerable growth in labour force participation by mothers of young children. In 1990, 36% of women with at least one child under six were employed. By 2002, Spanish mothers' employment rates had risen to 51% (OECD, 2005).

With the growth of mothers' employment, childcare issues have become an important matter for public concern. At the European Union level, the European Council of

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Barcelona (March 2002) agreed that "member States should remove disincentives to female labour force participation and strive (…) to provide childcare by 2010 to at least 90% of children between 3 years old and the mandatory school age and at least 33% of children under 3 years of age" (European Council, 2002). At the national level, the two major political parties concurring to general elections in 2008 (PSOE and PP) have included policies to increase the existing provision of childcare services in their electoral manifestos.

The perceived view is that, for mothers of preschool-age children, the decision to engage in paid employment typically implies the concurrent choice of a childcare arrangement. In this regard, the labour force participation of mothers of young children may exhibit sensitivity to the quality, the availability, and the cost of childcare (Han and Waldfogel, 2001, Viitanen, 2005).

The quality of childcare may be related to children's cognitive and non-cognitive outcomes.[1] Therefore, the quality of childcare may be a factor when parents decide whether to work and whether to use childcare. Even if there is no universal agreement as to what constitutes childcare quality (Han and Waldfogel, 2001), studies have mostly used childcare regulations (child-staff ratios, provider training, …) as proxies for quality. This literature yields mixed results, for some studies obtain no significant effects of quality on women's employment (Kimmel, 1998), and others obtain positive effects (Ribar, 1992, 1995).

Availability of childcare may also play an important role in women's employment decisions (Chevalier and Viitanen, 2002). Yet little empirical work has focused on the relationship between the geographic supply of childcare and female labour force participation. Most studies have included childcare availability measures as controls in childcare costs studies (Han and Waldfogel, 2001, Del Boca and Vuri, 2007). Those few studies that have focused on spatial accessibility have found positive effects of the geographic supply of childcare on women's labour participation in the U.S. (Herbst and Barnow, 2008, Stolzenberg and Waite, 1984), and no significant effects in Western Germany (Kreyenfeld and Hank, 2000).

Research on the relationship between childcare and labour market participation has mainly focused on the effect of childcare costs on women's employment decisions. In North America, most studies have estimated a discrete choice labour force participation probit with childcare costs and wages as key explanatory variables. Measures of expected childcare costs have been constructed, as these are usually available only for those who purchase childcare. These measures have been based on average cost in the community (Blau and Robbins, 1988) or selectivity corrected cost estimates (Connelly, 1992, Kimmel, 1998, Powell, 1997). This approach has also been common in Europe, particularly in the Netherlands (Van Gameren and Ooms, 2009, Wetzels, 2005), and some countries of the Mediterranean Europe (Del Boca, 1993, for Italy, and Dauli *et al.*, 2006, for Greece).

Other empirical studies for North America and Northern Europe have combined qualitative labour supply choices with childcare mode choices to form distinct combinations of labour supply and childcare that are estimated in a multinomial framework. Blau and Hagy (1998), Ribar (1995), Powell (2002), and Kornstad and Thoresen (2007), though following this approach, differed in the econometric estimation strategy they used. We may also find examples of this line of research for Continental Europe: Choné et al. (2003), for France, Lokshin (2004), for Russia., Lokshin and Fong (2006), for Rumania, or Whrolich (2006) for Germany.

In between these two approaches, some research has considered the interrelatedness of childcare choice and labour market behaviour but without incorporating choice of care mode (see, for instance, Cleveland et al., 1997, for Canada; Del Boca and Vuri, 2007, for Italy; and Viitanen, 2005, for the U.K.). These studies have usually estimated a bi-variate probit of labour force participation and paid childcare use, incorporating childcare costs and wages as key explanatory variables.

The empirical evidence gathered across these studies has generally supported the theoretical expectation that higher costs of childcare have a negative effect on the probability of participating in the labour market. The studies analyzing the Netherlands' situation constitute the exception, as childcare costs show no significant effect on Dutch mothers' labour force participation (Van Gameren and Ooms, 2009, Wetzels, 2005). Nonetheless, the range of elasticities has been large, ranging from –0.14 (Viitanen, 2005) to –0.98 (Connelly and Kimmel, 2003b), probably due to the different geographical scopes, methodologies used –probit, multinomial logit, bi-variate probit–, and sample characteristics –married/single mothers, age of youngest child,… However, as Herbst and Barnow (2008) state, there appears to be a recent convergence of estimates centering on –0.40.

In Spain these issues have been relatively neglected until very recently. The scarce literature has mainly focused on parents' time devoted to childcare from a household division of labour perspective as Fernández and Sevilla-Sanz (2006), García and Molina (1999), and Gutierrez-Domenech (2007). More recently, there have been some attempts to analyze childcare choice decisions by Spanish families, but without considering any consequences on the labour market (Borra and Palma, 2009, González and Vidal, 2006).

To our knowledge, just one other study has addressed the relationship between women's employment and childcare. Baizán and González (2007) use the Spanish Labour Force Survey (EPA) to analyze the effect of childcare availability on women's labour force participation. They do not control for the endogeneity of childcare use decisions, or for the effect of wages or childcare costs.

Data limitations may explain these restrictions and this relative lack of interest, as no single Spanish data set collects information on labour force status, childcare choices, and childcare characteristics (quality, availability and costs).

The aim of this paper is to provide evidence on the role that childcare costs play in the decision of mothers of preschool-age children to participate in the Spanish labour market. In order to overcome data limitations, we combine two different data sets: the Spanish Time-Use Survey (INE, 2002/2003) and the Spanish Household Budget Survey (INE 2005).

We are thus able to analyze both child care utilization and mothers' employment decisions, including reservation wages, (in the Time-Use Survey) along with child care costs (in the Household Budget Survey). To our knowledge, this is the first study that has examined the impact of childcare costs on female labour supply decisions in Spain. We also include and control for availability measures, but only at the regional level. We have not been able to incorporate any quality proxies.

In our study, we follow Viitanen (2005) and Del Boca and Vuri (2007) and simultaneously estimate the decision to become employed and the decision to use formal childcare. As robustness checks, we suggest different specifications to control for the mother's reservation wage and different sample selection criteria.

Our main empirical finding is that childcare costs exert a statistically significant and large negative impact on the decision to engage in paid employment. The estimated elasticity of labour force participation with respect to the hourly price of care ranges from –0.81 to –0.94. These figures lie within the upper end of the estimates found in previous literature, which, for Europe range from –0.14 in Viitanen's (2005) study for United Kingdom to –0.46 in Lokshin and Fong's (1998) study for Romania. Del Boca and Vuri (2007) do not report elasticities in their study for Italy. Nonetheless, our policy simulation results are quite similar to those found for Northern Italy.

The paper is organized as follows. Section 2 reviews the institutional setting in which Spanish families make their choices. Section 3 outlines the econometric model and estimation procedure issues. The data and variable construction are described in Section 4. Identification issues are discussed in Section 5. Section 6 presents empirical results and a discussion of the interpretation of the results and policy implications. The final section offers possible improvements and conclusions.

2. Institutional Setting

For the last two decades, Spain has witnessed a progressive accession of women to the labour market. Its female labour participation rates have risen about fifteen percentage points to exceed 61% in 2006, as shown in Table 1. Nevertheless, the figure is still weak compared to that of Northern European countries or the United States, which show participation rates of approximately 75%. A substantial part of this increase concerns the rise in the labour force participation rate of mothers. Even if in 1990 just 36% of mothers with at least one child under six were employed, by 2002, the employment rate of Spanish mothers had risen to 51% (Table 1).

Table 1
WOMEN'S LABOUR MARKET IN SELECTED OECD COUNTRIES

	Labour force participation rates		Employment rates for mothers with youngest child aged under 6	
	1990	**2006**	**1990**	**2006**
Australia	61,9	75,9	42,4	45,0
Belgium	46,1	65,9	64,4	68,8
Canada	69,2	77,9	56,9	62,7
Denmark	77,6	80,1	:	74,3
Finland	73,4	74,7	64,3	49,4
France	58,0	69,1	61,3	64,7
Germany	55,5	75,0	41,4	57,1
Greece	42,6	67,0	42,9	49,1
Ireland	42,6	71,3	30,6	51,8
Italy	44,0	62,7	45,3	53,0
Japan	60,4	73,1	37,2	35,2
Luxembourg	42,4	66,6	40,9	66,7
Netherlands	52,4	75,7	37,0	71,2
Portugal	59,6	73,9	67,4	79,2
Spain	**42,2**	**61,1**	**36,1**	**51,0**
Sweden	82,5	80,2	85,0	77,5
United Kingdom	67,3	76,7	42,5	57,0
United States	67,8	75,5	54,0	59,5
OECD	59,5	60,8	48,5	59,2

Source: OECD Employment Outlook 2007, OECD Society indicators 2005.

Nevertheless, Spanish women face a peculiar labour market. As stated by Saint-Paul (2000), among others, the labour market in Spain is subject to particular rigidities and imperfections, in particular concerning the structure of collective bargaining and employment protection legislation. These characteristics have tended to increase job security for full-time labour market participants, but at the cost of lower probabilities of employment for new entrants and/or individuals looking for part-time jobs. Table 2 suggests two possible consequences of these circumstances: a higher unemployment rate and a lower part-time employment rate. Unemployment rates in Spain have traditionally been the highest in Europe (Saint-Paul, 2000). In 2006, only Greece shows a higher female unemployment rate. Also, Spanish women have mostly full-time jobs. Part-time employment among Spanish women scarcely represents 21% of total employment. Together with Spain, countries as different as Greece, Portugal, Finland, Sweden, and United States present the lowest percentage of part-time workers. Nonetheless, as pointed out by Jaumotte (2003), a survey carried out in EU countries, which examined job preferences among couples with small children, found that only 11% Spanish families preferred the 'man full-time/woman part-time' option, compared to 67% in the Netherlands, 42.3% in Ireland, 41% in UK, or 28% in Italy. In the same vein, Ariza et al. (2003) note that in France, Greece, Portugal and Spain, part-time work is mainly involuntary, whereas in Denmark, Germany, Ireland, The Netherlands, and the UK, it can be considered the woman's choice.

Table 2
WOMEN'S UNEMPLOYMENT AND PART-TIME EMPLOYMENT IN SELECTED OECD COUNTRIES 2006

	Unemployment rate	Part-time employment as a proportion of total employment
Australia	5,1	40,7
Belgium	9,0	34,7
Canada	6,1	31,9
Denmark	4,6	25,6
Finland	8,1	14,9
France	10,7	22,9
Germany	10,3	39,2
Greece	13,5	12,9
Ireland	4,1	34,9
Italy	8,8	29,4
Japan	4,1	40,9
Luxembourg	5,8	27,2
Netherlands	4,8	59,7
Portugal	9,5	13,2
Spain	**11,6**	**21,4**
Sweden	7,2	19,0
United Kingdom	5,0	38,8
United States	4,7	17,8
OECD	6,6	26,4

Source: OECD Employment Outlook 2007, Jaumotte (2003).

Simultaneously, there has been a significant increase in the demand for non-parental care of preschoolers. These data are hard to obtain because of the different childcare arrangements, e.g. formal and informal, and because utilization rates for each type of childcare arrangement vary considerably with the age of the child.

Table 3 presents information from the OECD Family Database (2008), which brings together information from different OECD databases (such as, the OECD Social Expenditure database, the OECD Benefits and Wages database, and the OECD Education database).

As can be inferred from the second column, the situation for three-year-olds differs a great deal from one country to another. A partial explanation to this can be found in the different education laws. In Spain, at three years of age, children start what is called *Infant Education*, which precedes *Primary School*. And, even if it is not mandatory, public and private schools generally offer this cycle (3 to 5 years). The picture is not the same for children under three years of age. As the third column shows, in 2004, in Spain, only 21% of these children attended day-care centres or pre-schools. The figure is not too low, though, if compared to the less than 10% enrolment rates of Germany, Italy, and Greece. In fact, according to OECD's (2007b) report *Babies and Bosses*, on average across the OECD countries for which data are available, just 23% of zero- to three-year-olds use formal childcare.

Table 3
FORMAL* CHILDCARE IN SELECTED OECD COUNTRIES, 2004

	Enrolment rates for 3-years-old children	Enrolment rates for less than 3 3-years-old children	Childcare fee per two-year old attending formal care as percentage of average wage
Australia	55,0	29,0	22,4
Belgium	99,3	38,5	19,7
Canada	:	19,0	21,3
Denmark	81,8	61,7	8,4
Finland	37,7	22,4	7,6
France	100,0	26,0	25,1
Germany	69,5	9,0	9,1
Greece	:	7,0	4,5
Ireland	48,0	15,0	24,8
Italy	98,7	6,3	:
Japan	67,3	15,2	19,4
Luxembourg	37,9	14,0	32,4
Netherlands	32,3	29,5	17,5
Portugal	63,9	23,5	27,8
Spain	**95,9**	**20,7**	**30,3**
Sweden	82,5	39,5	4,5
United Kingdom	50,2	25,8	24,7
United States	41,8	29,5	19,5
OECD**	74,0	23,0	16,3

Source: OECD Family database (2008) and OECD (2007b) Babies and Bosses
* *Formal care refers to day-care centres and pre-schools.*
** *OECD averages across countries for which data are available.*

The affordability of purchased childcare services is mainly determined by the fee charged by providers. The last column in Table 3 shows average childcare fees per two-year old attending full-time (40 hours per week) formal childcare in different OECD countries. Fees are the gross amounts charged to parents, after any subsidies paid to the provider but before any childcare related cash benefits or tax advantages. The figures are calculated in relation to average gross wages (OECD, 2007b). As can be observed, formal childcare can be quite expensive in Spain. Only Luxemburg shows higher relative childcare costs. Together with these two countries, Portugal, France, Ireland, and United Kingdom exhibit childcare fees of at least approximately 25% of average wage.

In order to adequately take into account the time and money budget constraints faced by Spanish mothers, the above information must be completed with data on the financial support and time-related entitlements provided to parents. The second column in Table 4 provides information on statutory paid maternity leave for selected OECD countries. The entitlement to paid leave is presented as the full-time equivalent (FTE) of the proportion of the duration of leave if it were paid at 100% of wages –FTE=duration*wage replacement rate– (OECD, 2007b). The situation of Spanish mothers' can be considered slightly above average in this respect. Spain provides for 16 weeks of fully paid maternity leave around childbirth. Entitlement is conditional on previous work experience, though. There are also legal entitlements to paternity leave, for 15 days, fully paid, and parental leave, for three years, without income support.

Table 4
INSTITUTIONAL CONTEXT: FINANCIAL INCENTIVES AND LEAVE ENTITLEMENTS. 2005

	Full-time equivalent paid maternity leave	Family spending in cash and tax measures, in percentage of GDP
Australia	0,0	2,2
Belgium	11,5	2,2
Canada	8,3	1,0
Denmark	18,0	1,5
Finland	11,7	1,6
France	16,0	2,2
Germany	14,0	2,3
Greece	17,0	0,7
Ireland	14,4	2,3
Italy	16,8	0,6
Japan	8,4	0,8
Luxembourg	16,0	3,1
Netherlands	16,0	1,2
Portugal	17,0	0,9
Spain	**16,0**	**0,5**
Sweden	12,0	1,5
United Kingdom	12,0	2,2
United States	0,0	0,8
OECD	13,5	1,5

Source: OECD Family database (2008).

Together with childcare provision and leave entitlements, a third way to deal with the reconciliation of work and family life is to grant financial aid to families with young children. The third column in Table 4 offers family spending in cash and tax measures as percentage of GDP for selected OECD countries. Total expenditures include cash transfers –either income-related or universal child allowances and leave payments– and financial support delivered through the tax system. Spain, spending just 0.5% of GDP, ranks very low, only followed by Greece and Italy. Even if cash transfers constitute the dominant component of financial support directed towards families, in Spain, child allowances are income-related, with relatively low income thresholds, what determines that only 5% of households are eligible for this benefit –compared to 16% in Italy, 73% in France or 100% in Northern Europe– (Del Boca *et al.,* 2009). Financial support delivered through the tax system has a more general incidence, especially for families with children under three years old. Apart from general child tax allowances, Spain grants an additional tax allowance for each child under three and also a refundable tax credit for working mothers with a child under three. Due to these provisions Spain ranks highest in De Henau's *et al.,* (2006) appraisal of child-related financial support of European Union countries.[2]

Given this institutional background, we can conclude that Spain shares many features of the 'pro-traditional' model with other Southern European countries, Italy, Portugal, and Greece (Del Boca *et al.,* 2009). In this model, the main concern is the preservation of the family, governments taking very limited responsibility for supporting it. Coincident with the 'male breadwinner model' it encompasses, in Spain young children's responsibility and care rely on their mothers. They may decide to remain in the labour market after giving birth, in which case, non-parental care is generally needed. Usual arrangements are, in order of importance: day care centres, care by relatives, schools, and baby-sitters. Nevertheless, as emphasized by De Henau *et al.* (2006), Spain also distinguishes itself from other Southern European countries. Spain's reliance on child-related financial support, especially for children under 3, together with its relative lack of public childcare facilities, encourages market care. And in this respect Spain is similar to UK (De Henau *et al.,* 2006).

Given the above discussion, in this paper we will study the work-childcare options of Spanish mothers with children under 3 years of age, that is, children not eligible for *Infant Education.*

3. Empirical Model

The behavioural model underlying the empirical work in this paper follows the work of Ribar (1995), Blau and Hagy (1998), or Del Boca and Vuri (2007), and a brief verbal description follows. Women are assumed to be the principal caregivers in the household and, therefore the employment decisions of family members other than the mother are taken as given. We also suppose two forms of childcare: informal care provided by the mother, father, or other household members and formal, paid care, purchased in the market. Mothers are assumed to maximize utility, where utility is expressed as a function of leisure time, market goods, and childcare quality.

The constraints in this maximization problem include constraints on the mother's and child's time, a money budget constraint and a production function for childcare quality. The maximization of this utility function subject to the constraints yields the mother's demand for leisure (labour supply) and her demand for paid childcare services. The functions can be specified as:

$$H^* = \alpha_L + \beta_L W + \chi_L P_C + \delta_L X_L + \varepsilon_L \quad (1)$$

$$C^* = \alpha_C + \beta_C W + \chi_C P_C + \delta_C X_C + \varepsilon_C \quad (2)$$

where H^* is mother's time in market work; C^* is child's time in paid childcare; W is the hourly wage faced by the mother; P_C is the price of childcare; X_L is a vector of other determinants of the decision to engage in paid employment as age, non-labour income, household composition,…; X_C is a vector of other usual determinants of the decision to purchase paid childcare as age of the child, availability of alternative care arrangements, presence of other children,…; and finally, ε_L and ε_C are the error terms.

In these equations, the key parameter of interest is χ_L, the effect of childcare costs on female labour supply. In a context like the one outlined above, characterized by significant labour market rigidities, shifts in individual choice are likely to take place at the extensive margin (McFadden, 1974): that is, whether to work or not to work. Therefore, following Del Boca and Vuri (2007), Han and Waldfogel (2001), Viitanen (2005) and Wetzel (2005), among others, we will focus on the dichotomous labour force participation (LFP) and not the continuous labour supply decision. H^* in Equation (1) can be considered a latent variable, and LFP, its dichotomous counterpart (LFP = 1 if $H^*>0$, 0 otherwise). Also, childcare may involve non-negligible fixed costs, resulting from a minimum-hours requirement or travel time expenses (Joesh and Hiedemann, 2002, Ribar, 1995). So in order to adequately model labour force participation decisions, we have to take into account childcare use (CCU) decisions. In this case C^* in Equation (2) is the latent variable, and CCU, its dichotomous counterpart (CCU = 1 if $C^*>0$, 0 otherwise).

The error terms in Equations (1) and (2) are likely to be correlated. The literature on employment and child care issues discussed above emphasizes that the decision to enter the labour force and the decision to use of formal childcare are interrelated decisions. That is, if a mother's preferences for work are related in unobservable ways to her preferences for childcare, then the choice of her work status will have to be modelled simultaneously with her childcare decision (Hallman *et al.,* 2002). Therefore, following Cleveland *et al.* (1996), Del Bocca and Vuri (2007), and Viitanen (2005), we jointly estimate these equations by means of a bi-variate probit model, assuming ε_L and ε_C are jointly normally distributed with mean 0, variance 1 and covariance ρ (Heckman, 1978).

Apart from the simplest case in which both equations are modelled as independent probits (with $\rho = 0$), econometric alternatives to this model include multinomial logit and nested logit. None of them seems to be adequate. Choices in the multinomial logit have to fulfil the independence of irrelevant alternatives assumption by which the conditional probability between any pair of choices has to be independent of other alternatives. It is not very likely that the odds ratio between the option 'working and using formal childcare' and 'not working and using formal care' is independent of the option 'working and not using formal care', for instance. In the nested logit model a sequential model structure has to be assumed. And as stressed by Kornstad and Thoresen (2007), in modelling work status and childcare choices, it is not obvious which dimension, childcare or employment, is decided first. Therefore, in this paper, the more general simultaneous framework of the bi-variate probit model is used.

Following Kimmel (1998), the hourly wage W and the hourly price of care P_C are entered in the equations as two distinct terms because the total number of hours worked per week is not constrained to be equal to the number of paid childcare hours. In other words, the model allows mothers to purchase more or less hours of childcare than their working hours and even using childcare when they are not working. See below the discussion on the endogenous variables for the Spanish data.

Prior to estimating Equations (1) and (2), we have to solve the issue of the likely endogeneity of wages and childcare costs. Current wages may be endogenous because they also reflect time allocation decisions, that is, they are only observed for currently working mothers (Gupta and Stratton, forthcoming). In order to circumvent this endogeneity problem, we estimate predicted wages $\hat{W}$ potentially faced by all women in the sample. Using the standard approach first suggested by Heckman (1976), a reduced-form participation probit is first estimated across all women in the sample. The probit parameters from this participation equation are used to generate the inverse Mills ratio and thereby take into account the possibility of sample-selection bias in the wage equation. The wage equation is estimated using the sample of workers only. The results of the wage equation are then used to generate values of the predicted wage for all women in the sample (Rammohan and Whelan, 2007, Viitanen, 2005).

Similarly, childcare costs may also be endogenous. We only observe childcare costs paid by the household and this constitutes a measure not only of price, but also of utilization (Del Boca and Vuri, 2007). To get around this likely problem, we use predicted prices $\hat{P}_C$ potentially paid by all women in the sample. Following once more Heckman's (1976) methodology, a reduced form probit on the use of formal childcare is first estimated. The estimates from the childcare use probit can then be used to construct the sample selection term. A cost of childcare equation is then estimated over the sample of childcare users with the inclusion of the sample selection term. And the estimates from the equation modelling childcare costs are used to compute the predicted price of childcare for all women in the sample.

Most studies using North-American data employ a double selection model, since in many datasets childcare costs are only observed for households where the mother is employed. Therefore in addition to the selection regarding utilization, employment selection is also controlled for (Cleveland *et al.,* 1997, Kimmel, 1998, Powell, 1997). However, in Spain, as in Italy (Del Boca and Vuri, 2006) or Germany (Wrohlich, 2004), the link between employment and childcare use is not so strong (see Table 6), and, datasets provide information on childcare costs for non-working mothers too. Therefore, a single sample selection correction term is employed, as in Del Boca and Vuri (2006), Wrohlich (2004) and Viitanen (2005).[3]

4. Data and Variable Construction

The study uses primarily data from the Spanish Time-Use Survey-STUS (INE, 2002/2003).[4] The STUS is part of the Harmonized European Time-Use Surveys (HETUS) launched by the EU Statistics Office (Eurostat). Technically it is a household-based survey with multiple questionnaire components conducted in 2002-2003. For our study, information contained in the household and individual questionnaires has been used.[5] Even if it is not specifically intended for studying childcare and labour supply, the survey provides interesting information on households' childcare arrangements and on the employment status of household members.

For our study, out of the 20,603 households, 1,970 households for which the youngest child was less than four years old and non-eligible for *Infant Education* were initially selected. Lack of data on critical variables, together with difficulties in determining primary childcare type[6] led to a reduction in sample size of 342 observations. Dropping those observations had virtually no impact on the characteristics of our sample. Because the focus of this study is on the elasticity of female LFP to the price of formal childcare, children primarily cared for at public schools or by babysitters were also excluded. This selection led to a minor reduction in the average age of the child and a slight decline in the proportion of high educated women in the sample. Finally, we also excluded observations using primarily relative care. This final selection obviously resulted in an increase in the proportion of children in formal care. Nevertheless, in order to account for the potential bias originating from this decision, as a robustness check, we included these observations along with the parental, no-paid care option.[7] Table 5 presents the characteristics of our initial and final samples, as well as the effects of the successive selections described above.

Table 5
EFFECT OF THE SAMPLE SELECTION ON SELECTED VARIABLES

	All households with children under 3	Cleaned sample	Final sample with relative care	Final sample without relative care
Sample size	1970	1628	1347	1078
Child in formal childcare	27,93%	28,05%	30,86%	41,37%
Mother in paid employment	44,21%	42,57%	44,45%	41,00%
Age of the child	1,45	1,40	1,24	1,28
Age of the mother	33,50	33,38	33,06	33,16
Mothers educational level				
Primary or less	41,57	42,33	45,34	45,54
Secondary	32,46	32,15	31,65	32,37
University	25,96	25,50	23,00	22,07

Source: Spanish Time-Use Survey, INE 2002/2003.

As already stated, our endogenous variables are LFP and CCU. Construction of the first variable is straightforward. LFP is coded 1 if the mother reports positive hours of work during the previous week (INE, 2004a). As regards childcare utilization, the household questionnaire asked families whether each of their children under ten were taken care of by different alternatives and for how long (in weekly hours) these different arrangements took place.[8] A family is coded as using childcare services if they reported using paid formal childcare for the longest number of hours per week, compared to care by relatives or babysitters and care at schools.[9] Non-using families are those for which no such regular external care was recorded, either in paid childcare, by relatives or babysitters, or at schools (Blau and Hagy, 1998). Nonetheless, in section 6.2, as a robustness check, we include observations on families whose youngest child is regularly cared by relatives along with the latter non-using families.

Table 6 provides a simple tabulation of our endogenous variables. Out of the 1078 total, 442 mothers or 41.0 % are employed and 446 or 41.3% report using formal childcare. Although we will consider these issues in detail later, we would like to underline two facts. The first one is that approximately 25% of the surveyed non-working mothers use paid, formal care for their children. This fact has also been mentioned by Del Boca and Vuri (2007), for Italy, or Wrohlich (2006), for Germany. In the UK this fraction is only about 5% (Viitanen, 2005). The second is that a non-negligible 36% of working mothers rely exclusively on parental care or care provided by any adult member living in the household.[10]

Table 6
FORMAL CHILD-CARE USE AND LABOUR FORCE PARTICIPATION

	Non-working	Working	Total
Non-Using	474	158	632
	(74.53)	(35.74)	(58.62)
Using	162	284	446
	(25.47)	(64.25)	(41.37)
Total	636	442	1.078
	(100.00)	(100.00)	(100.00)

Source: Spanish Time-Use Survey, INE 2002/2003.

Additionally, the STUS contains detailed information on the income, labour market activities, socio-demographic characteristics of the household and its members, and the autonomous region and municipality size of the city of residence of the family. Unfortunately this database does not provide information on the expenditure involved in the activities, and thus the price of childcare, our key explanatory variable, could not be computed. Thus, information from other sources had to be collected. Specifically, we used data on the Spanish Household Budget Survey (INE, 2005) for the same years (2002-2003).

The Spanish Household Budget Survey (INE, 2005) provides detailed information on expenditures incurred by families in different seven digit COICOP/HBS[11] categories, together with data on household income and information on the region and size of the municipality where the family resides. Following Del Boca et al. (2005), the above two data sets were merged using propensity-score matching methods (see Appendix B for details in the procedure). The aim of this method is to match an individual of the Time-Use Survey with a similar individual of the Household Budget Survey according to some chosen criteria in order to collect relevant information from both surveys. Specifically, to calculate day care prices, we imputed Kindergarten Expenditures (1231208-COICOP-HBS) for an individual of the Time-Use Survey using the information available from a similar individual from the Household Budget Survey.

Descriptive statistics for the variables used as controls in the estimations, together with estimated prices for users, are reported in Table 7. Our most important control variable is hourly wages. As stressed by Amuedo-Dorantes and de la Rica (2009), one drawback of the STUS is that the information on net monthly earnings is reported in intervals of 500€, instead of as a continuous variable. This may create measurement errors, particularly at the upper tail of the earnings distribution. Fortunately, only 0.2% of our sample falls within the last interval. Following Amuedo-Dorantes and de la Rica (2009) and Gutierrez-Domenech (2008) we have used class marks to approximate earnings values within each class. Wages of working mothers are then computed by dividing estimated monthly earnings by reported weekly hours worked times the average number of weeks per month. Given the relevance of this variable, we have tried a different specification in section 6.3 as a robustness check.

Table 7
DEFINITION AND BASIC STATISTICS OF DEMOGRAPHIC AND SOCIOECONOMIC VARIABLES

	Units	Definition	Mean
PRICE	Eu/hour	Hourly price of childcare of users	1,071 (0.23)
WAGE	Eu/ hour	Hourly market wage of workers	6,844 (5.85)
AGE	years	Age of the child in years	1,288 (1.01)
AGE_0	0/1	Dichotomous variable which takes value 1 if the child is less than one years old.	0,277 (0.44)
AGE_1	0/1	Dichotomous variable which takes value 1 if the child is less than two years old.	0,294 (0.45)
AGE_2	0/1	Dichotomous variable which takes value 1 if the child is two or three years old.	0,42 (0.49)
AGEMOTH	Years	Age of the mother	33,145 (5.30)
EDLEVEL1	0/1	Dichotomous variable which takes value 1 if the mother's education level is primary school or less	0,454 (0.49)
EDLEVEL2	0/1	Dichotomous variable which takes value 1 if the mother's education level is secondary school	0,322 (0.46)
EDLEVEL3	0/1	Dichotomous variable which takes value 1 if the mother's education level is University degree	0,223 (0.41)
FOREIGNER	0/1	Dichotomous variable which takes value 1 if the mother is a foreign person	0,083 (0.27)
ONE-PARENT	0/1	Dichotomous variable which takes value 1 if the mother is single	0,019 (0.13)
CHILDREN	number	Number of children under 10 living in the household	1,862 (0.92)
ADCHILDREN	0/1	Dichotomous variable which takes value 1 if there are additional children under 10 living in the household	0,615 (0.48)
ADULTS	number	Number of adults living in the household	2,089 (0.35)
UNINCOME	Thou.eu/month	Aggregated monthly earnings of household members less mother's labour income	1,338 (0.90)
LESS-TENTH	0/1	Dichotomous variable which takes value 1 for municipalities with population under 10.000	0.111 (0.38)
AVAILABILITY	Places/child	Regional availability of day-care places per child	0,041 (0.02)
CARE_WAGE	Thou.eu/year	Regional average wage of workers in the personal services sector	11,347 (1.37)
UNEMPLOYM	Percentage	Regional unemployment rate	17,185 (7.16)

Source: Spanish Time-Use Survey (INE 2002/2003), Spanish Household Budget Survey (INE 2003), Anuario de Estadísticas Laborales y Asuntos Sociales. 2003 (Ministerio de Trabajo y Asuntos Sociales, 2004), Encuesta de Estructura Salarial. 2002 (INE, 2004) and Encuesta de Población Activa, Resultados Anuales. 2003 (INE, 2004)

Other control variables include characteristics of the child (his age), characteristics of the mother (age, education level, foreign status,…) and characteristics of the household (number of adults, presence of additional children, number of children,…). An important economic feature of the family is non-labour income defined as aggregated monthly earnings of household members less mother's labour earnings. Again we have followed Amuedo-Dorantes and de la Rica (2009) and Gutierrez-Domènech (2008) in using class marks to approximate earnings values within each class for both total household income and total mother's labour earnings, so that non-labour income could be approximated.

The final data set was completed by adding regional information on availability of childcare places from the Anuario de Estadísticas Laborales y Asuntos Sociales (Ministerio de Trabajo y Asuntos Sociales, 2004), average wage rates of women working in the Personal Services Sector from the Encuesta de Estructura Salarial (INE, 2004b), and regional unemployment levels from the Encuesta de Población Activa, Resultados Anuales (INE, 2004c). A description of these variables is also provided in Table 7.

5. Identification

As stated by Kimmel and Connelly (2007), estimating multistep models such as this requires strict attention to equation identification. We confront these issues at two levels: estimation of the supporting equations and estimation of the labour participation/childcare use bi-variate probit model.

At the first level, the wage and hourly childcare cost are estimated by Heckman (1976) procedures. Technically, this type of model can achieve identification by functional form assumptions (Cameron and Trivedi, 2005). In practice, nonetheless, most researchers feel more comfortable if at least one regressor in the dichotomous index equation is excluded from the continuous outcome equation. Variables incorporated in the reduced form labour participation equation that are not included in the wage equation include non-labour income, number of children in the household, number of adults, if the child is an infant, and availability of formal childcare. All of them have been used in previous studies (Cleveland *et al.,* 1996, Connellly and Kimmel, 2003a, Kimmel, 1998, Kimmel and Connelly, 2007, Viitanen, 2005, van Gameren and Ooms, 2009). These variables directly affect the mother's reservation wage, and hence her employment decision, without determining her wages. As discussed below in section 6, estimation shows that non-labour income and the number of children are significant predictors of labour participation for Spanish mothers.

Variables included in the reduced-form childcare use equation that are not entered in the childcare price equation include the presence of additional children, the number of adults in the household, the fact of living in a small town, and the regional availability of childcare places. As the number of children in a family unit grow, taking care of the children at home and not working results in a cost saving strategy, as opposed to working and placing the children in childcare. Also, mothers living in households with other adults are likely to choose home care for their children. Neither of these circumstances is likely to affect formal childcare prices. The regional proportion of children attending day-care and the size of the municipality are considered proxies of childcare availability. We expect availability to be a good candidate for an identifier too, predicting the probability of paying for childcare, but not affecting the amount paid.[12] Most former studies have used family composition variables as identifiers of the selection term (Cleveland *et al.,* 1997, Kimmel and Connelly, 2007, Rahmohan and Whelan, 2007, Viitanen, 2005, Whrolich, 2004). Only Del Boca and Vuri (2007) and van Gameren and Ooms (2009) include also availability indicators as identifiers. Estimation shows that the fact of living in a small town is a significant predictor of Spanish mothers' childcare use.

The second level of concern with identification comes from the use of these predicted regressors in our bi-variate probit model of labour participation and childcare use. As indicated by Connelly and Kimmel (2003a), what is needed to identify the price of childcare and wage variables are variables included in those estimating equations that are excluded from the final bi-variate probit. Again we look for exclusion restrictions that can both be justified theoretically and have empirical significance in the first stage. Following Connelly and Kimmel (2003a), van Gameren and Ooms (2009) and Wetzels (2005), among others, we use the age of the mother and education levels as identifiers. The education variables satisfy the criteria of empirical significance in the first-stage equations. The theoretical argument is that education does not directly affect employment and childcare use, but through the effect of estimated wages and prices for care. To aid identification, regional dummies are also excluded from our bi-variate model as in Cleveland *et al.* (1996).

6. Empirical Results

In this section, we discuss the results from estimating the labour participation/childcare use bi-variate probit. We also report elasticities, perform robustness checks, and explore public policy implications. Results from the supporting equations for wages and childcare costs are presented in Appendix C.

6.1. Bi-variate Model Results

Estimated coefficients for the primary LFP/CCU bi-variate probit equations are given in Table 8. The regressors in this equation include the predicted hourly wage and the predicted hourly price of childcare from Tables C.1 and C.2 in Appendix C, along with other socio economic characteristics of the household. Wages have a significant positive effect on both labour force participation and paid childcare use, while the hourly cost of childcare shows a negative significant impact on both decisions. In addition, the estimated correlation coefficient (rho) is positive and significant, indicating the adequacy of the simultaneous estimation of both equations. These results are all consistent with the underlying behavioural model.

Table 8
LFP/CCU BI-VARIATE PROBIT COEFFICIENT ESTIMATES

	Corr. Coef. Rho:	0.595***	Log-likehood	–1.149.674
		(0.041)	Chi2(18)	368.320
	Chi2(1)	133.283	Prob > chi2:	0.000
	Prob > chi2:	0.000		

	LPF		CCU	
Variable	**Coef.**	**Boots. S.E.**	**Coef.**	**Boots. S.E.**
CONSTANT	0.167	0.572	1.366**	0.680
PRICEHAT	–0.697**	0.292	–0.723*	0.411
WAGEHAT	0.324***	0.038	0.148***	0.038
AGE_0	–0.207*	0.110	–1.342***	0.147
AGE_1			–0.390***	0.128
FOREIGNER	–0.139	0.159	–0.154	0.178
ONE_PARENT	–0.225	0.400	0.080	0.422
UNINCOME	–0.286***	0.056	0.048	0.050
AD_CHILDREN	–0.344***	0.081	–0.236**	0.091
ADULTS	0.304**	0.131	–0.284**	0.139
AVAILABILITY	–1.639	2.187	2.551	2.069
LESS_TENTH			0.403***	0.131
UNEMPLOYM	–0.024***	0.007		
CAPITAL	0.129*	0.079		

*Significance level: * 10%; ** 5%; ***1%.*
Standing errors are computed by bootstrapping with 200 repetitions.

Controlling for childcare costs, the presence of additional children in the household continues to have a significant negative impact on female LFP, as also reported by Del Boca and Vuri (2007), for Italy, Daouli et al. (2004), for Greece or Powell (1997), for Canada. Consistent with the expected income effect, as found by Cleveland et al. (1997), for Canada, Daouli et al. (2004), or Viitanen (2005), for the UK, though in opposition to del Boca and Vuri (2007), higher levels of income earned by family members other than the mother are found to affect labour participation decisions negatively. As reported by Powell (1997) and Cleveland et al. (1996), though contrary to Viitanen's (2005) findings, when both wages and childcare costs are controlled for, the mother's immigrant status does not significantly affect her labour participation decision. On the contrary, labour market conditions, included through the regional unemployment rate, are still significant determinants of female labour participation, as in Del Boca and Vuri (2007).

As previously found by van Gameren and Ooms (2009), for the Netherlands, and Kreyenfeld and Hank (2000), for western Germany, the regional provision of formal care has no significant effect on labour force participation decisions. This contradicts previous evidence for Spain, as Baizán and Gonzalez (2007) report a significant positive influence of childcare availability on female employment.[13]

With respect to the childcare use decision, one of the most significant determinants continues to be the age of the child, with older children being more likely to be cared for at day-care centres (see Viitanen, 2005). Once the hourly price of childcare and the mother's expected wage are controlled for, availability of informal modes of care, measured by the presence of adults in the household, has a significant negative effect on the probability of using market care, as also found in Cleveland et al. (1996), del Boca and Vuri (2007) and Viitanen (2005). Additionally, compared to families with only one child, mothers are much less likely to rely on purchased care if they have more than one child under the age of 10. Cleveland et al. (1996) and van Gameren and Ooms report similar results.

Surprisingly, the regional availability rates of day-care places do not significantly affect the probability of using paid, formal care. This is in contrast with results by Chiuri (2000), for Italy, or van Gameren and Ooms (2009). However, given that the positive sign of this variable is intuitively correct, we feel that more disaggregated data may have resulted in more accurate estimates.

6.2. Elasticities

Participation and childcare use elasticities based on the estimation results in this paper are reported in Table 9. Our main empirical finding is that the expected price of childcare exerts a statistically significant and large negative impact on the decision to engage in paid employment. The elasticity of labour force participation with respect to the hourly price of care is –0.93, indicating that reducing childcare costs by 10% would lead to a 9% increase in the probability of engaging in paid employment. This figure lies within the upper end of the estimates found in previous literature which range from –0.14 in Viitanen's (2005) study for United Kingdom to -0.98 in Connelly and Kimmel's (2003b) study for single mothers in the USA.[14] Former studies for Southern European countries (del Boca and Vuri, 2007, Dauli *et al.,* 2004) do not report elasticities.

We may speculate on the likely sources of this large elasticity. Connelly and Kimmel (2003a) contend that studies, such as this one, that rely on predicting childcare prices from individual characteristics tend to get larger elasticities than studies that rely on regional childcare price data. In their view, one of the most important aspects of the market for childcare is that individuals face different costs for similar services depending on the availability of low- or no-cost childcare options and only individual based models, such as ours, can take this variation into account. In addition, as we have been able to see in the Institutional Setting section, Spanish families face considerably larger childcare costs as a fraction of average wages than other OECD families (OECD, 2008). As the Slutsky equation shows, the greater the proportion of income allocated for a commodity, the more elastic will its demand be.

The elasticity of labour force participation with respect to the mother's wage is 0.96. Previous estimates are quite similar (0.81 for Cleveland et al. (1996), 0.85 for Powell (1997)), with the exception of 3.25 in Kimmel's (1998) analysis of the USA, 0.35 in Kornstad and Thoresen's (2007) study of Norway, and 0.42 in Viitanen's (2005) analysis of United Kingdom.

The elasticity of paid childcare use with respect to its own price is –0.98. This indicates that a 10% reduction is childcare costs would increase the probability of using market care by about 10%. This figure lies within the range of former estimates, which vary from –0.46 for United Kingdom (Viitanen, 2005), to –1.06 for Canada (Cleveland *et al.,* 1996), or –1.86 for the United States (Ribar, 1992).

Finally, the elasticity of paid childcare use with respect to the mother's wage is 0.45, indicating that high-wage mothers are more likely to purchase formal childcare. In particular, a 10% increase in the mother's wage rate is associated with a 4% increase in the probability of using paid childcare. This elasticity is also within the range of previous estimates (i.e., 0.18 in Cleveland *et al.,* 1997, or 0.62 in Viitanen, 2005).

Table 9
PRICE AND WAGE ELASTICITIES FROM LFP/CCU MODEL. 1078

	LPF		CCU	
	Elasticity	**S.E.**	**Elasticity**	**S.E.**
PRICEHAT	–0.930**	0.388	–0.981*	0.556
WAGEHAT	0.961***	0.120	0.447***	0.117

*Significance level: * 10%; ** 5%; ***1%.*

6.3. Robustness Checks

This section analyzes sensitivity of the previous results to different specifications and samples.

6.3.1. Indirect consideration of wages

Computation of working mothers' hourly wages might be subject to measurement errors, due to the interval nature of the earnings variable. In order to account for this potential problem, we check whether our results are robust to a different specification in which wages are indirectly taken into account. We estimate a model in which the reservation wage is instrumented by the mother's educational level and age, as in Joesch (1998).[15] Women with more education typically earn a higher wage rate than those with less education. Thus the higher the educational level, the higher the probability of working and using childcare services.[16] The age of the mother is considered an indicator of paid work experience. More experience translates into better pay and, thus, an incentive to stay in the work force when a child is born, ceteris paribus.

Table 10 shows price elasticities from a bi-variate probit model including the mother's age and educational level as regressors, but not estimated wages.[17] The figures remain virtually unchanged suggesting that the main results are not subject to potential measurement bias.

Table 10
ROBUSTNESS CHECK I. PRICE ELASTICITIES W/O WAGEHAT. 1078 Obs

	LPF		CCU	
	Elasticity	**S.E.**	**Elasticity**	**S.E.**
PRICEHAT	–0.948**	0.359	–0,961*	0.554

*Significance level: * 10%; ** 5%; ***1%.*

6.3.2. Inclusion of relative care

In order to test the sensitivity of our results with respect to the sample selection criteria being used, we estimate equations (1) and (2) including the 269 observations pertaining to the relative care choice, along with the parental care, no paid care option. This specification involves assuming that families do not distinguish between taking care of preschool children at home by household members or relying in other relatives living in a different household. Table 11 shows the elasticities computed using this new sample. As can be observed, both wage and price elasticities are slightly inferior. In fact, the estimated elasticity of labour force participation with respect to the hourly price of childcare indicates that a 10% reduction of childcare costs would increase the labour participation rate of mothers of pre-school-age children by approximately 8%. This estimate can be considered a floor –and the former 9.3%, a ceiling– for the actual effect of a childcare price reduction on mothers' labour participation rates.

Table 11
ROBUSTNESS CHECK II. PRICE AND WAGE ELASTICITIES. 1351 Obs

	LPF		CCU	
	Elasticity	**S.E.**	**Elasticity**	**S.E.**
PRICEHAT	–0.812**	0.315	–0.867***	0.334
WAGEHAT	0.904***	0.100	0.268***	0.082

*Significance level: * 10%; ** 5%; ***1%.*

6.4. Policy Implications

Finally, to discuss some public policy implications of our estimates, we simulate the employment effects of different levels of childcare costs subsidization. Specifically, we have calculated the mean predicted probabilities of labour force participation for direct childcare subsidies of 25%, 50% and 100%. The subsidy simulation provides estimates of the anticipated degree of mothers' employment response in the event of significant childcare subsidy. The results of these simulations are shown in Table 12, together with results from similar exercises.

The mean predicted probability for our original sample is 39.9%. This measure is very close to the actual participation rate in the sample, which is 41.0%. If childcare costs were subsidized by 50%, the model predicts a LFP rate of 58.9%. If childcare costs were fully subsidized, mothers' LFP would rise to 75.9%. These simulations indicate that the LFP of Spanish mothers is quite responsive to subsidized childcare. Furthermore, when families using primarily relative care are included in the sample, the changes in LFP rates predicted by the model remain quite similar. In particular, universal childcare subsidization implies that 79.5% of Spanish mothers would be employed in this case. These results are similar to those found for Italian mothers living in non-rationed areas (Del Bocca and Vuri, 2006). Studies for North America show slightly reduced effects.

Table 12
LABOUR FORCE PARTICIPATION SIMULATIONS

Study	Country	Baseline	25% sub.	50% sub.	100% sub.
This study	Spain	39.90%	9.50%	18.90%	35.90%
This study with relative care	Spain	45.60%	9.30%	18.30%	33.90%
Del Bocca and Vuri (2006)	Italy (North)	61.50%		15.50%	26.50%
	(South)	40.80%	:	2.70%	5.40%
Kimmel (1998)	USA	58.00%		5.00%	9.00%
Powel (1997)	Canada	46.40%	:	9.50%	16.80%
Connelly (1992)	USA	58.80%	:	5.20%	9.90%

7. Conclusions

This paper has analyzed the effect of childcare costs on the labour supply decision of Spanish mothers using primarily data from the Spanish Time Use Survey. This is done through the estimation of bi-variate probits on the probability of using formal paid childcare and the probability of engaging in paid employment on the labour market -both decisions being functions of expected childcare costs, expected wages and other household characteristics. In order to circumvent the potential endogeneity of both wages and childcare costs, sample-selection corrected estimates of expected costs and wages are used.

The key finding in this paper is that childcare prices significantly affect the labour force participation of Spanish mothers. This result is robust to different specifications and sample selection criteria. A commonly argued rationale for government subsidization of childcare costs is to raise the labour force participation of mothers. The responsiveness of the labour supply of mothers to childcare costs estimated herein indicates that such subsidies do encourage female labour supply.

We would like to emphasize that quality and availability are also important. In our analysis, regional availability rates were not found significant, but we feel that better, more disaggregated data could result in more accurate predictions.[18] Lack of individual data on childcare quality is also recognized as a limitation of the paper. Analyzing childcare in Spain is complicated by the fact that many of the relevant factors in childcare choices, such as distance to close relatives, the number and type of childcare services within a convenient distance, child-staff ratios, or flexibility in the hours of service offered, are unobserved in the available data. More detailed information would allow researchers to control for any supply-side restrictions when analyzing the impact of childcare on mothers' labour force participation.

Notes

1. There is limited consensus in the literature about the effects of childcare quality on children outcomes, though. For disadvantaged children, the literature suggests that participation in high-quality programs aimed directly at this group –what the U.S. literature refers to as 'early childhood intervention' (Waldfogel, 2002)– is beneficial to participating children (Currie, 2001, NICHD and Duncan, 2003, Peisner-Feinberg et al., 2001). However, there is little convincing evidence that the quality of universal childcare has a positive effect on child development (Blau and Currie, 2006). Blau (1999) finds insignificant effects for the U.S. while McMahan (1992) reports positive effects for France.
2. De Henau et al. (2006) compute the financial support indicator comparing total child benefit packages (related to the presence of a child in the household) received by a family with one child aged under six to those obtained by a family with one child aged at least six years.
3. As a control, we estimated predicted hourly childcare costs in a model with double selection, but the correlation between the childcare cost equation and the labour participation equation was found to be non-significant. The results are available from the author upon request.
4. Other research using this data set can be found in Amuedo-Dorantes and de la Rica (2009). Fernandez and Sevilla-Sanz (2006), Gimenez et al. (2007) or Gutiérrez-Domenech (2007).
5. The survey also contained information on daily activities by means of the completion of a personal diary for each individual aged 10 or older.
6. See the discussion on the construction of the CCU variable below.
7. See the robustness check in section 6.3.
8. See Appendix A for further details.
9. The vast majority of families used only one childcare mode. In the data cleaning stage, we omitted 15 cases in which we could not decide which childcare mode was used longest. Nevertheless, including those observations together with 7 additional cases which reported using day care centres, but for a shorter period than relatives or schools, had almost no impact on our findings (Results available upon request).
10. This explains why we decided to include the category relative care along with no care use in our sensitivity analysis of section 6.3.
11. The Classification of Individual Consumption by Purpose Adapted to the Needs of Household Budget Surveys (COICOP-HBS) is an international coding system designed for household budget surveys implemented in many countries (INE 2005).
12. Our specification includes regional dummies which may capture differences in regional economic development which may affect both the likelihood of using formal childcare and the price paid for it.
13. Nevertheless, note that the authors use provincial, instead of regional, data.
14. Blau and Robbins (1988) obtain -0.38, Cleveland et al. (1996), -0.39, Ribar (1992), -0.74 and Lokshin and Fong (2006), -0.46. The studies analyzing the Netherlands show no significant effect of childcare costs on Dutch mothers' labour force participation (Van Gameren and Ooms, 2009, Wetzels, 2005)
15. In our estimation of wages, the educational level appeared as the most important predictor of potential wages.
16. In fact, Gupta and Stratton (forthcoming) favour education-based measures over earnings-based measures that need correction for sample selection.

17. Identification of the childcare price equation is now based on regional dummies only.
18. County level data or data relative to Spanish provinces would be very illuminating.
19. The contents of this section rely heavily on Borra and Palma (2009).
20. Comparison available upon request.
21. Table B.1 describes the variables finally used in the matching process together with their logit estimates.
22. Within each block, we tested for equality of means of each characteristic between the treated and the control units. This is a necessary condition for the balancing property. The results are not reported but are available upon request.
23. The standard errors are bootstrapped.

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Resumen

La participación laboral femenina en España es de las más bajas de Europa. Este trabajo analiza el grado en que la participación laboral femenina se ve afectada por los costes de la atención formal a la infancia. Ambas decisiones, la participación en el mercado de trabajo y el uso de servicios formales de atención, se consideran de forma conjunta por medio de un modelo probit bi-variante que considera la selección muestral. Con información procedente de la Encuesta de Empleo del Tiempo español, el estudio indica que la participación laboral de las madres españolas es muy elástica a los cambios en los costes de atención a la infancia.

Palabras clave: costes de cuidado de niños, participación laboral femenina.

Clasificación JEL: J13, J22, C35

Appendix A. Spanish Time-Use Survey Questions

Relevant questions from the Spanish Time-Use Survey used in the construction of the CCU variable (INE, 2004a):

2. For household members aged under 10 only

State any care received and/or school status of children aged under 10 years, starting with the youngest and in increasing order of age, and the average weekly hours of such childcare.
(You may indicate more than one option if needed.)

For each child by increasing order of age:
Household member sequence number: _ _

Regular childcare
Cared for by individuals:
1. Relatives in the household
*2. **Unpaid** persons (relatives outside the household, friends, neighbours, etc.).*
3. Paid persons (babysitters, child care specialists, etc.)
(If Yes, record average weekly hours)

Attends an institution
4. Kindergarten, infant school, crèche
5. School
(If Yes, record average weekly hours -include hours devoted to extracurricular activities within the institution (provided that they are regular-).

If the child attends an institution, please fill out the following details:
6. The institution where they spend the most time is: Public / Private
7. Do the children have lunch at the institution? Yes / No

Appendix B. Statistical matching

In this section we explain how the statistical matching was performed.[20] Matching involves pairing units, from different datasets, that are similar in terms of their observable characteristics. In this case, the recipient dataset was the Spanish Time Use Survey-STUS (INE 2002/3) and the donor dataset was the Spanish Household Budget Survey-SHBS (INE 2005), both representative of the Spanish population. These two sets of information share important observable characteristics.

As Dehejia and Wahba (2002) stated, with a small number of characteristics (for example, two binary variables), matching is straightforward (one would group units in four cells). However, when there are many variables, as in the present situation, we need to define a function which measures the similarity between the individuals of the two samples and solves the dimensionality problem. As emphasized by Del Boca *et al.* (2005), propensity score-matching methods (Rosenbaum and Rubin, 1983) provide specific criteria to assign to each individual of the recipient data set a similar 'individual' from the donor dataset. Each pair of individuals created according to this procedure will give origin to an integrated record, with the relevant information from both surveys.

For the present study, 402 observations from the SHBS were used along with the 446 observations from the STUS. As a baseline analysis, we compared the averages for the variables the two surveys had in common and checked that apparently both surveys were very similar for childcare services users.[20]

Next we needed to match observations from the two surveys. As first suggested by Rosenbaum and Rubin (1983), we used the conditional probability of belonging to one of the samples (the so-called propensity score) to reduce the dimensionality of the matching problem previously discussed. This propensity score was computed as $p(X_i) = Pr(i \in T.U.|X_i = x)$ (Del Boca *et al.,* 2005). Therefore, rather than match on the regressors, matching was performed on $p(X_i)$.

Table B1
PROPENSITY SCORE COEFFICIENT ESTIMATES

	Number of obs: 848	Log-likelihood –577.08528 LR chi2(31) 101.95 Prob > chi2: 0.000	
	Definition	**Coef.**	**Std. Err.**
CAPITAL	Capital cities	0.7450***	0.25
THOUSAND	Municipalities with population over 100.000	0.9213**	0.36
FIFTY	Municipal. with population betw. 50.000 & 100.000	1.8193***	0.34
TWENTY	Municipal. with population betw. 20.000 & 50.000	0.333	0.30
TENTHOU	Municipal. with population betw. 10.000 & 20.000	1.0254***	0.29
LESTENTH	Municipalities with population under 10.000	Ref.	
MONOPAREN	= 1 if the mother is single	1.2141**	0.52
EMPLOYEE	=1 if the mother is an employee	–0.3150*	0.17
CHILDREN	Number of children under 10 in the household	0.1623	0.10
ADULTS	Number of adults living in the household	0.3953*	0.23
INCOME	Aggregated monthly earnings of household members	0.0001	0.00
IN_DOMSTAFF	=1 if the household has live-in domestic staff	–1.1695	1.32
HOTWATER	=1 if the dwelling has hot water	–0.3566	0.29
LANDPHONE	=1 if the dwelling has landline phone	–0.2995	0.25
TENURE	=1 if the dwelling is owned	0.3392*	0.18
SECONDHOME	=1 if the household uses a second home	0.3803	0.26
_cons		–1.3103*	0.72

*Significance level: * 10%; ** 5%; ***1%.*
Specification includes regional dummies.

In order to compute the propensity score, we run a logit regression of the binary indicator taking value 1 for observations in the Time-Use sample (and 0 for the Household Budget sample) over the set of common household characteristics. We followed the algorithm proposed by Becker and Ichino (2002), which tests if the propensity score satisfies the balancing property (see Cameron and Trivedi 2005).[21] We ended up with five blocks; in each of them the score was balanced across the treated units and controls.[22]

As Del Boca *et al.* (2005) indicate, since the propensity score is a continuous variable, exact matches will rarely be achieved and a certain distance between individuals belonging to the two samples has to be allowed. Thus, we chose to use kernel-based matching (Heckman et al. 1998), where we associate a kernel-weighted average of the outcome of all donor-dataset units to the unit *i* of the recipient dataset:

$$\hat{y}_i = \frac{\sum_{j \in H.B.} K\left(p_i - p_j\right) y_j}{\sum_{j \in H.B.} K\left(p_i - p_j\right)} \tag{3}$$

where $K(u) \propto exp(-u^2/2)$ is Gaussian, p_i and p_j are the propensity scores of units $i \in T.U.$, and $j \in T.U.$, and y_j stands for the outcome of individual $j \in T.U.$ As can be seen, the weight given to donor unit *j* is in proportion to the closeness between *i* and *j*.

After the statistical matching was performed, each individual from the STUS using paid childcare services was imputed the expenditures of a 'similar' individual from the SHBS.

Finally we proceeded with an internal evaluation of the statistical matching. We compared the average values of the imputed variable after the matching and the corresponding average in the donor set, that is, the SHBS sample. This difference was 1.7% and not significant at conventional levels of testing.

Appendix C. Estimating Wages and Childcare Costs

Table C.1 presents a selectivity corrected log-wage model of the mother, where the selection refers to the decision of engaging in paid employment. The results are consistent with those usually found in the labour supply literature. As reported for example by Powell (1997), increases in the mother's level of education and age have a significant positive effect on both participation and wages. Also, on average, immigrant mothers present lower participation rates and receive lower wages. As found by Viitanen (2005), the number of children under age ten is associated with decreased female labour participation. Regional unemployment rates, included to control for labour demand conditions, have the expected negative effect on both participation and wages (Kimmel, 1998). Finally, household non-labour income is used to identify the model as it has a direct effect on the mother's reservation wage, hence affecting her employment decision, with no impact on her wage.

Non-labour income has the expected negative effect on the employment probability (Viitanen, 2005). Consistent with model expectations, the sample selection term shows a significant positive impact, indicating that working mothers tend to obtain higher wages than non-working mothers.

Table C1
LFP PROBIT COEFFICIENT AND LOG-WAGE COEFFICIENT ESTIMATES

Number of obs:	1078	Log-likehood	–754.475
Censored obs	636	Chi2(7)	145.360
Uncensored obs	442	Prob > chi2:	0.000

	LPF		Log-wage	
Variable	**Coef.**	**S.E.**	**Coef.**	**S.E.**
CONSTANT	–0.176	0.400	0.602***	0.225
AGEMOTH	0.003	0.009	0.013**	0.006
EDLEVEL2	0.594***	0.101	0.390***	0.078
EDLEVEL3	1.124***	0.116	00.886***	0.089
FOREIGNER	–0.853***	0.186	–0.599***	0.153
UNEMPLOYM	–0.038***	0.007	–0.016***	0.005
UNINCOME	–0.216***	0.047		
CHILDREN	–0.100**	0.046		
ADULTS	0.123	0.100		
AGE_0	–0.105	0.080		
AVAILAB	–0.019	1.725		
LAMBDA			0.520***	00.059

*Significance level: * 10%; ** 5%; ***1%.*
Specification includes regional dummies.

Results of the selectivity corrected log-childcare costs model are shown in Table C.2.[23] The age of the child and the level of education of the mother have the expected impact on the use of childcare. As found in Powell (1997), having older children significantly increases the likelihood of paying for care. Also, as reported by Viitanen (2005), more educated mothers are more likely to purchase childcare. Surprisingly, the presence of adults or other children under ten years of age in the household does not significantly affect the probability of using formal childcare, once other household and family characteristics are controlled for.

As expected, the age of the child is a significant determinant of childcare prices. The regional wage rate, included to control for supply conditions, is significant and of expected sign. Contrary to intuition, the educational level of the mother is negatively related to childcare costs. Nonetheless, it should be kept in mind that less educated mothers are likely to use less hours of care, as they are probably not working, and possibly face higher hourly prices. Regional dummies, not shown for brevity, are also quite significant, indicating the importance of regional variation in childcare costs. The coefficient on the selection term is negative and significant. This result suggests that families purchasing childcare face lower prices than non-users.

Table C2
CCU PROBIT COEFFICIENT AND LOG-PRICE COEFFICIENT ESTIMATES

Number of obs:	1078	Log-likehood	–519.626
Censored obs	632	Chi2(23)	115.520
Uncensored obs	446	Prob > chi2:	0.000

	CCU		Log-Price	
Variable	**Coef.**	**Boots. S.E.**	**Coef.**	**Boots. S.E.**
CONSTANT	–1.177**	0.513	–0.401*	0.226
AGE	0.577***	0.058	–0.047***	0.004
AGEMOTH	–0.007	0.007	–0.001	0.003
EDLEVEL2	0.343***	0.095	–0.053	0.047
EDLEVEL3	0.662***	0.107	–0.123**	0.061
ONE_PARENT	0.450	0.355	–0.210**	0.096
FOREIGNER	–0.363**	0.155	0.017	0.060
UNINCOME	0.047	0.058	–0.005	0.013
CARE_WAVE	0.039	0.032	0.059***	0.018
AD_CHILDREN	–0.121	0.088		
ADULTS	–0.153	0.147		
LESS_TENTH	–0.494***	0.154		
AVAILABILITY	1.361	2.053		
LAMBDA			–0.258***	0.02

*Significance level: * 10%; ** 5%; ***1%.*
Specification includes regional dummies.