Entry into Motherhood: The effect of wages

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Abstract

Using the ECHP, we explored the determinants of having the first child in Spain. Our main goal was to study the relation between female wages and the decision to enter motherhood. Since the offered wage of non-working women is not observed, we estimated it and imputed a potential wage to each woman (working and non-working). This potential wage enabled us to investigate the effect of wages (the opportunity cost of time non-worked and dedicated to children) on the decision to have the first child, for both workers and non-workers. Contrary to previous results, we found that female wages are positively related to the likelihood of having the first child. This result suggests that the income effect overcomes the substitution effect when non-participants opportunity cost is also taken into account.

1. INTRODUCTION

Since the New Home Economics (Becker, 1965; Willis, 1973) appeared, many articles have emphasized the effect of wages and income on fertility (Butz and Ward, 1979; Hotz and Miller, 1988; Heckman and Walker, 1990; among others). In particular, most papers have found a negative relationship between female wages and fertility. This result seems to support the suggestion that the incorporation of women into the labour market (encouraged by increasing wages) is closely connected to the decline in fertility, since the evidence points to a clear positive relationship between female wages and female participation rates (Becker, 1965; Killingsworth and Heckman, 1986). The negative effect of female wages on fertility rates is mainly due to the negative relationship beween female participation and fertility rates (Willis, 1973; Becker and Tomes, 1976).

To disentangle the effect of wages on fertility, net of their effect through participation, we need to observe the wage offered to non-participant women, i.e. we need to observe the wage offered that is turned down. As this is not possible, we impute a wage (estimated potential wage) to all women based on their characteristics and estimate its effect on the probability of having the first child in Spain.

In this paper we studied the probability of having a first child by estimating a reduced form model using the ECHP data for Spain. We imputed wages for all women, both workers and nonworkers to measure the effect of wages on the probability of entering motherhood. What we were therefore trying to compute was a measure of the opportunity costs of women in the decision to have the first child. Although we can observe this value for participants simply by observing their wages, we cannot observe it for non-participants, despite the fact that these women can take into account the price of their time in the market when taking the fertility decision. We estimated the probability of entering motherhood using a discrete hazard approach that enabled us to employ time varying variables using the method proposed by Jenkins (1995). Then we

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estimated both decisions (participation and having the first child) simultaneously to determine the effect of wages on the first child probability net of the effect on the participation decision. We found a positive wage effect on the probability of having the first child, when using wages for all women. The positive effect stands when modelling fertility and female participation simultaneously. This result does not mean that the neoclassical model is not supported by our work, as we also found a negative relationship between participation and fertility in our simultaneous estimation. It means instead that the income effect predominates when using wages for all women, and that given their participation decision, women with higher earnings from work are more likely to have their first child in Spain. Our result provides support for the public policies subsidizing working mothers introduced by several European governments.

The paper is structured as follows. Section 2 presents the stylized facts and previous study results for Spain. Section 3 presents the estimation methods applied to study the effect of female wages on the probability of having the first child. Section 4 describes the dataset used in our work as well as the variables controlled for and presents how we have estimated the potential wage for all working and non-working women. Section 5 presents the empirical results. In Section 6 we do a simulation exercise to show the effect of a public policy that subsidizes births for working women. Section 7 presents the principal conclusions.

2. STYLIZED FACTS AND PREVIOUS STUDIES ON SPAIN

Although the fall in the Total Fertility Rate came later than in other OECD countries, the Spanish TFR soon became one of the lowest in the world (Figure 1). At the same time, Female Participation Rates continued to increase, mainly among middle-aged women (Figure 2). However, Female Participation Rates in Spain are still lower than in other European countries with higher Fertility Rates. One of the reasons some researchers have found to explain this difference (Ahn and Mira, 2001) is the country's high unemployment rate (Figure 3). Adverse labour market conditions may have a major effect on a woman's decisions about acquiring human capital. Unemployment in Spain has been especially high among women and young people. This may have created an incentive to increase the time dedicated to education for two reasons: for one thing, the opportunity cost of studying decreases if the probability of young people finding a job is low. Second, the probability of finding a job increases with education (Nickell, 1997). With regard to fertility decisions, more time spent in education delays motherhood, as few women have children while at school (in Spain the number of children per women in education is 0.04). Adverse labour market situations are analyzed in Adsera (2004), who found that high unemployment and job instability in Southern Europe helped to explain lower fertility rates in comparison with northern European countries. De la Rica and Iza (2005) found that fixed-term contracts delayed entry into motherhood in Spain. Adsera (2005) found that unemployment and temporary contracts reduced the number of children. The age of entry into motherhood in Spain rose (figure 4) from 24.2 years in 1976 to 29.2 in 2003. Postponing the age at which women have their first child influences the drop in the Total Fertility Rate (Ryder, 1964; Boongarts and Feney, 1998). Ortega and Kohler (2001) estimated that the Spanish TFR net of the calendar effect was slightly higher than the period TFR, with the completed fertility of the cohort born in 1960 being around 1.6 children per woman (Bernardi and Requena, 2003). However, postponing the age at which women produce their first child narrows the span of time in which women are exposed to maternity and may reduce the final parity[2].

[2] The difference between desired parity and actual parity in Spain is the highest of all European countries, at almost 1 child (Delgado and Castro, 1998).

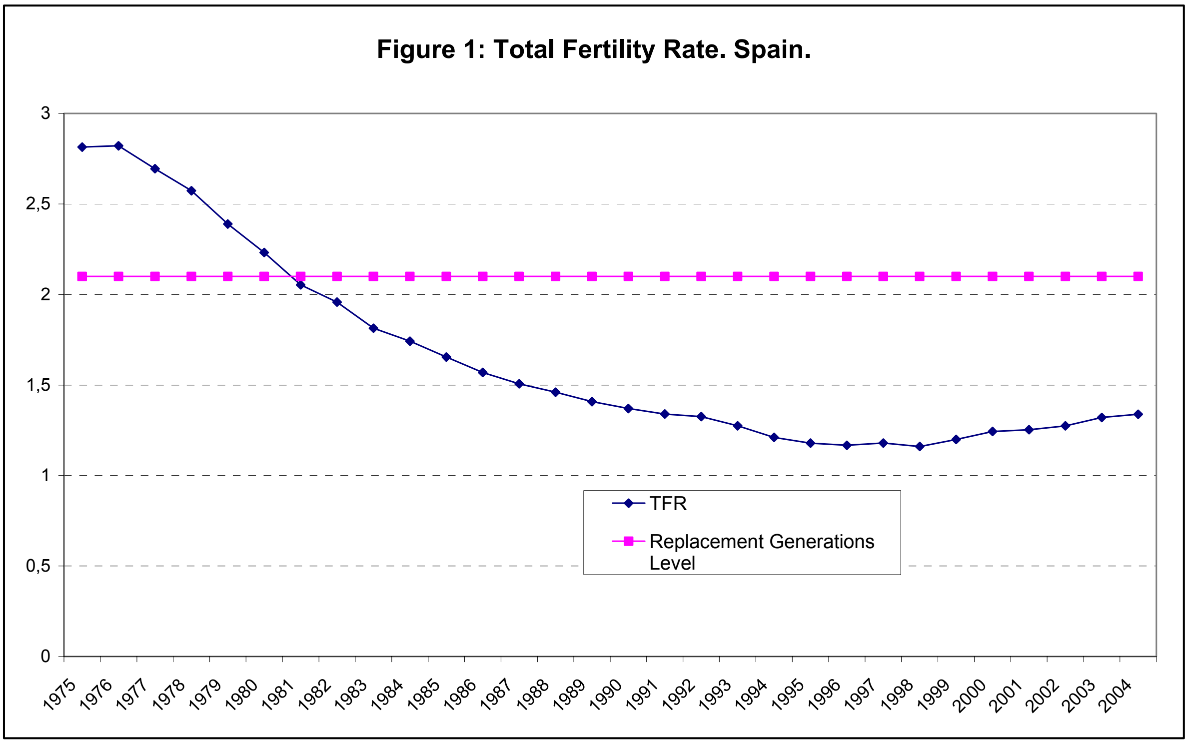
Figure 1: Total Fertility Rate. Spain.
3
2,5
2
1,5
1
0,5
0
TFR
Replacement Generations Level
1975
1976
1977
1978
1979
1980
1981
1982
1983
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1987
1988
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Figure 2: Female Participation Rates. Spain.

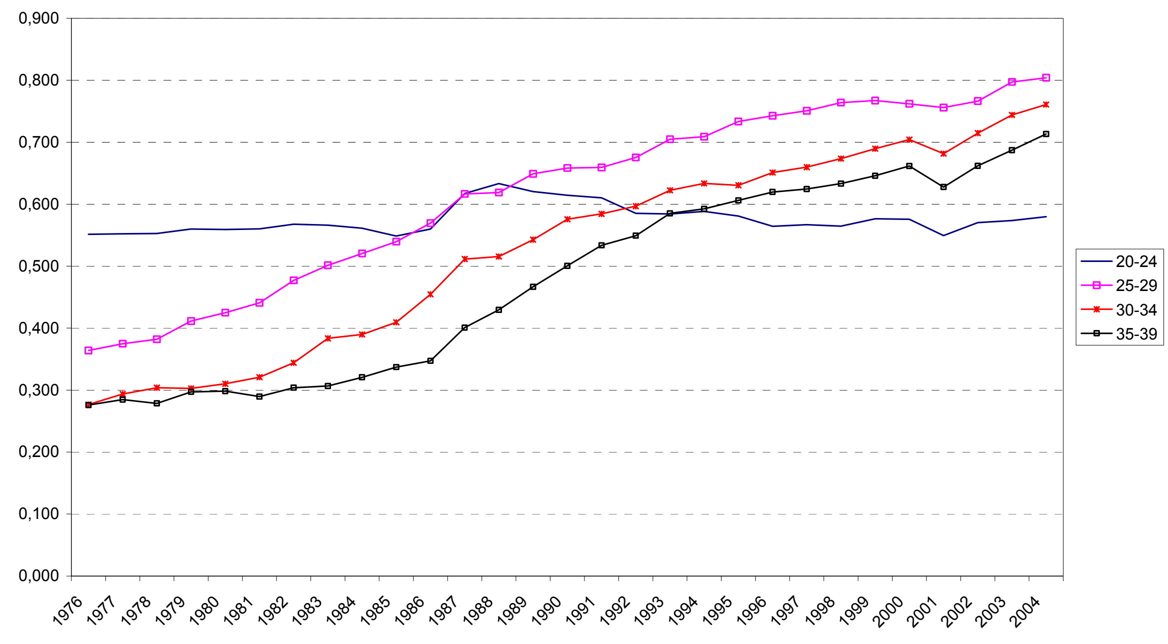


Figure 3: Female Unemployment Rates. Spain.

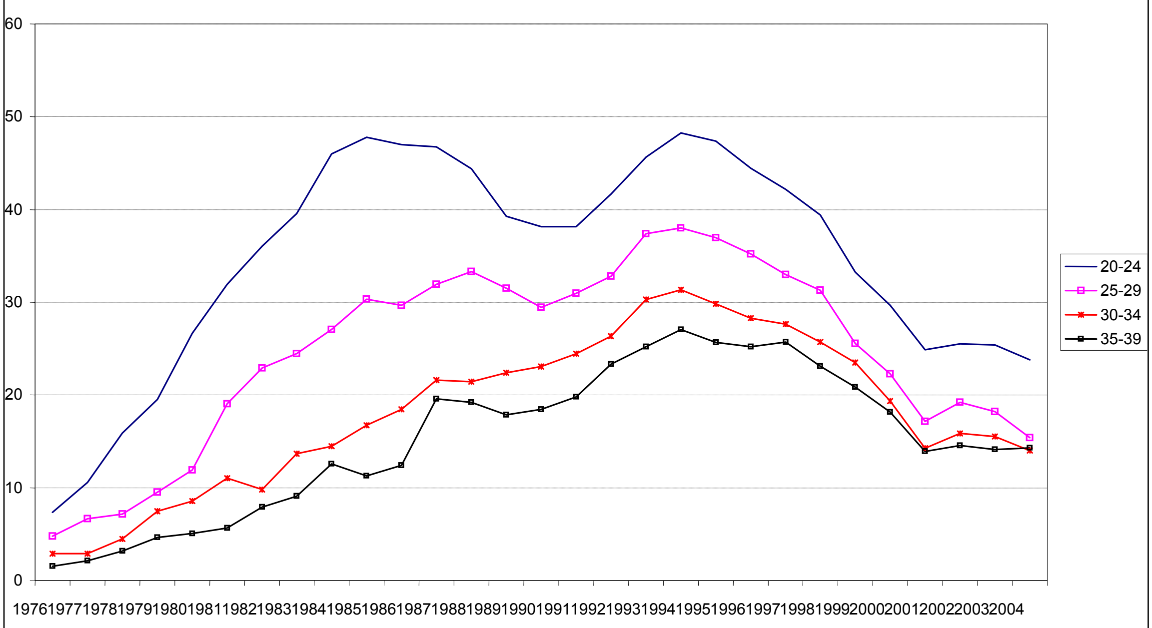
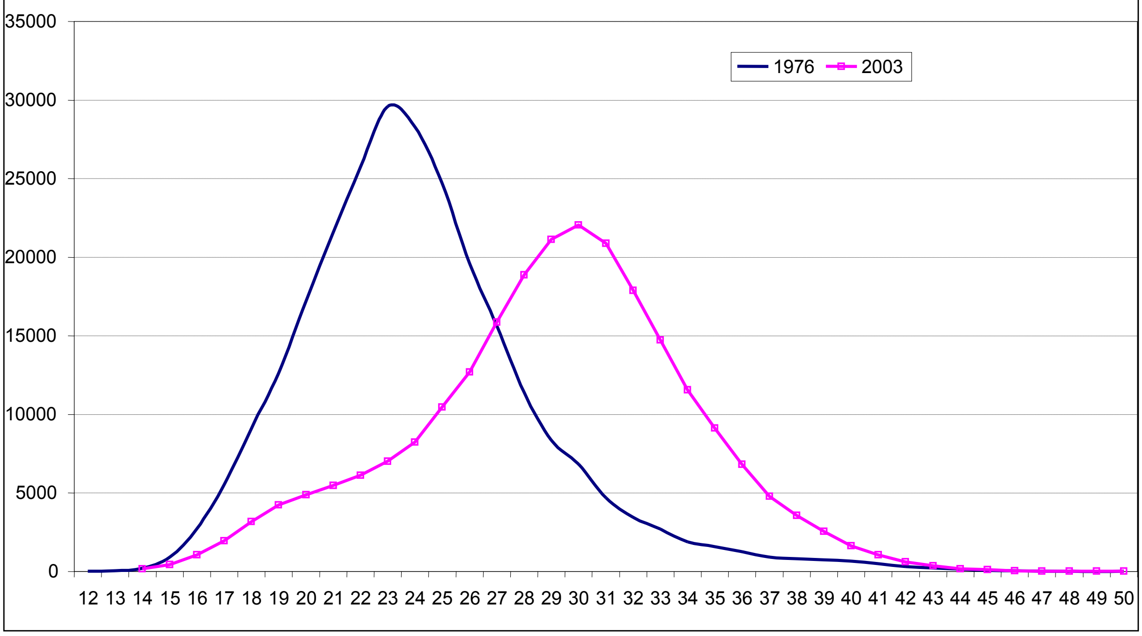


Figure 4: First Births by Mother's Age, Spain.



3. ESTIMATION METHODS

The two econometric models used to estimate the effect of female wages on the probability of having a first child are described below. They are the discrete hazard model and the bivariate model. We began by estimating the probability of having the first child without explicitly considering the endogeneity between fertility and labour participation decisions (discrete hazard model), and we compared the results with an estimation in which we took account of the participation and birth decisions simultaneously (bivariate model).

3.1 Discrete hazard model

Given the discrete nature of our data, we applied the methodology developed by Allison (1982) and Jenkins (1995) to study the probability of having a first child. A reduced form model was considered. Our sample consisted mainly of young women (46% of the women in the sample belonged to the 18-24 age group) who were more likely not to have had children at the interview date. This class of sample selection in duration models is well-known as *stock sampling* (Lancaster, 1990). Jenkins (1995) explains how to deal, in a very simple way, with a discrete duration model that faces a stock sampling type of sample selection, and so we used his methodology.

The hazard rate, i.e. the probability of having the first child at t conditional on remaining childless until t, in discrete time is:

$$h_{it} = prob\left(T_i = t \middle| T_i \geq t; X_{it}\right) \qquad (1)$$

where X_{it} is a vector which includes independent covariates, some of them time-varying, and T_i is a discrete variable that measures the time lapsed until the first child's birth (measured in quarters). Jenkins (1995) shows that the non-conditional probabilities can be expressed as:

$$prob\left(T_i = t\right) = \frac{h_{it}}{(1-h_{it})} \prod_{k=1}^{t} (1-h_{ik}) \qquad (2)$$

and

$$prob\left(T_i > t\right) = \prod_{k=1}^{t} (1-h_{ik}) \qquad (3)$$

where t = 1 is the first moment at which the woman is able to conceive a child. Typically, this period is asumed to be the period in which the woman reaches age 15. However, rather than observing the woman from this age, we observed her from a period we called r, being r > 1.

The non-conditional probability that a woman may conceive at any period s (s >1) would be:

$$(1-h_{is})\,(1-h_{is-1})(1-h_{is-2})\,...(1-h_{i1}) \qquad (4)$$

But we observed women since r (r>1), therefore, the probability that a woman may conceive at any period s (s >r) conditional on having remained childless until r would be:

$$\frac{(1-h_{is})(1-h_{is-1})(1-h_{is-2})\ldots(1-h_{i1})}{(1-h_{ir-1})(1-h_{ir-2})\ldots(1-h_{i1})}=(1-h_{is})(1-h_{is-1})(1-h_{is-2})\ldots(1-h_{ir}) \quad (5)$$

Thus, the stock sampling type of sample selection can be easily handled by term cancellation and probability is readily definable (see Jenkins, 1995). The probability for individual i may be written as:

$$L_i=\left[h_{it}\prod_{t=1}^{d_i-1}\left[1-h_{it}\right]^{c_i}\right]\left[\prod_{t=1}^{d_i}\left[1-h_{it}\right]^{(1-c_i)}\right] \quad (6)$$

where d_i is the number of periods the individual remains in the sample.The value of d_i is 28 for women who do not give birth during the whole period of observation (censored spells) and, for women who have a child, the value of d_i is the number of lapsed quarters from the time they are first observed until they have the child (completed spells). Thus, c_i is an indicator that takes value one for women with complete spells and 0 otherwise. Redefining our sample by creating as many observations as the periods a woman may conceive and defining a dummy y_{it} that takes value 0 for each of the quarters in which conception is possible and 1 for the exit quarter, the individual likelihood function is:

$$L_i=\prod_{t=1}^{d_i}\left[\frac{h_{it}}{1-h_{it}}\right]^{y_{it}}\prod_{t=1}^{d_i}\left[1-h_{it}\right] \quad (7)$$

The estimation of this model is equivalent to the estimation of a cross-section model with binary dependent variable (logit or probit) in the redefined sample as noted above. In particular, using a complementary specification log-log one for the hazard, the model converges to a proportional hazard model as the rate of exit becomes smaller. Following Heckman and Singer (1984), we control for unobserved heterogeneity in a semi-parametric way. We assume that unobserved heterogeneity follows a discrete distribution function with two mass points (types). For a model with types z = 1,2; the hazard function for an individual belonging to type z, is:

$$h_{zt} = 1 - \exp\left[-\exp\left\{\beta' X_{it} + \beta_0 + m_z\right\}\right] \qquad (8)$$

and the probability of belonging to type z is also estimated.

If both unobserved heterogeneity and duration dependence are insignificant, the model is equivalent to estimating a logit model with the expanded sample (woman/quarter observations).

3.2 Bivariate Probit

One of the main problems in trying to measure the influence of the female labour market situation on fertility is that both variables are potentially endogenous (see for example Browning, 1992). The consensus between researchers is that both decisions are taken jointly. A woman plans her family and professional life simultaneously. Thus, the estimations of direct influences of fertility on female labour supply and vice versa when this simultaneity is not taken into account suffer from endogeneity bias. A standard remedy for this problem is the instrumental variables technique, which consists of finding variables correlated to fertility but not to female participation. However, instrumental variables techniques are not valid when both variables (fertility and female participation) are discrete. This is the case when measuring the relation between the participation decision and the decision to have the first child. Carrasco (1998) uses the sex of the first two children as an instrument since it is related to the decision to have a third child but not to the participation decision. In the estimation, she uses a switching probits model where the switching variable (fertility) is treated as endogenous. Using data from the Labor Force Survey, Alvarez (2001) uses this technique to measure the effect of the birth of a child on the mother's participation decision for Spain. Nakamura and Nakamura (1992) point out that the search for valid instruments may be fruitless.

To model both fertility and participation decisions avoiding endogeneity we considered a bivariate probit model. The two equations that we specified and estimated simultaneously, were the fertility status at a particular quarter and the participation status one year before potentially having the first child. This model is:

$$\begin{aligned} F_i &= x_i^{'}\beta + u_{1i} \\ P_i &= z_i^{'}\beta + u_{2i} \end{aligned} \qquad (9)$$

where

$$F_i = \begin{cases} 1 & \textit{if the woman gives birth} \\ 0 & \textit{otherwise} \end{cases} \qquad (10.1)$$

and

$$P_i = \begin{cases} 1 & \textit{if the woman participates} \\ 0 & \textit{otherwise} \end{cases} \qquad (10.2)$$

and u_{1i} and u_{2i} follow a Bivariate Normal distribution:

$$\begin{pmatrix} u_{1i} \\ u_{2i} \end{pmatrix} \sim N\left(\begin{pmatrix} 0 \\ 0 \end{pmatrix} \begin{pmatrix} 1 & \rho \\ \rho & 1 \end{pmatrix} \right) \qquad (11)$$

where ρ is the correlation between both error terms. Thus the log-likelihood for this model is:

$$L(\pi / Y_i) = \prod_{i=1}^{n} \pi_{00}^{Y_i(0,0)} \pi_{10}^{Y_i(1,0)} \pi_{01}^{Y_i(0,1)} \pi_{11}^{Y_i(1,1)} \qquad (12)$$

Where $Y_i = (F_i, P_i)$ is an indicator function and

$$\begin{aligned} \pi_{00} &= \int_{-\infty}^{0} \int_{-\infty}^{0} \Phi_2 \left(u_{1i}, u_{2i} / \rho\right) du_{1i} du_{2i} \\ \pi_{10} &= \int_{0}^{\infty} \int_{-\infty}^{0} \Phi_2 \left(u_{1i}, u_{2i} / \rho\right) du_{1i} du_{2i} \qquad (13) \\ \pi_{01} &= \int_{-\infty}^{0} \int_{0}^{\infty} \Phi_2 \left(u_{1i}, u_{2i} / \rho\right) du_{1i} du_{2i} \\ \pi_{11} &= 1 - \pi_{00} - \pi_{01} - \pi_{10} \end{aligned}$$

Φ_2 being the bivariate normal cumulative distribution function.

4. THE DATABASE

4.1 The Survey

In the present paper we used data for Spain from the European Community Household Panel (1994-2001). The ECHP collects demographic and labour information of all the members of the households interviewed (7206 households in Spain in 1994). Thus, we had the age, month of birth, marital status, labour situation and wages of all members of the household. The labour information is available monthly for the entire year previous to the interview, so we can analyze the situation of the woman at the moment of the decision to have the first child, which we assume is taken a year before the birth. The main advantage of this survey in respect of other Spanish surveys is the availability of detailed information on income and wages. This information enables us to estimate a measure of the opportunity cost of the decision not to work (wage) for both types of women, participants and non-participants in the labour market (the way this opportunity cost is estimated is described in the Appendix). The major drawback of the ECHP is the short panel span, which prevents us from observing the complete fertility history of the women and relating it to the labour situation. Instead of monthly information, we used quarterly information (quarter as time unit), largely because of the low number of exits occurring in the sample. We considered a reduced form model in which we studied the determinants of the probability of having a first child, focusing on the effect of the female wage on this probability

4.2 Final Sample and variables

After excluding from the sample the women who did not provide all the information needed, we were left with a sample of 2292 women aged 17-35 who had not had their first child when first interviewed. Unfortunately, the lack of information on the type of labour contract held by the worker in the previous year (information was only available on the current type of labor contract) would have made us lose a wave to be able to use this information, so our final sample had 1792 women aged 18-34, the first time they entered the sample, observed over an average of 10 quarters (from a minimum of only 1 quarter to a maximum of 28 quarters). Thus, our expanded sample had 26490 women/quarter observations. Out of the total number of women in the sample, 327 (5.5%) had their first child in the observed period of time. We defined our dependent variable as a binary variable taking value 1 if the woman has the birth in the quarter and 0 otherwise. We related the result of this variable to individual and household characteristics. All variables are delayed a year to consider the time of gestation of the child, i.e. to consider the time between the actual birth and the decision to have a first child, assumed to be four quarters earlier. Some individual characteristics, such as the labour situation, may change from one quarter to the next, but we assumed they did not vary within each quarter.

Table 1 shows descriptive statistics of our sample. Average age of women in our sample is 25.1 years and they were observed over a maximum of 28 quarters (7 years). As we have already mentioned above, Jenkins' method to estimate a discrete duration model is equivalent to estimating a binary dependent variable model (logit or probit or its extensions) using the expanded sample (each woman contributes as many times to the sample as the periods in which she may conceive) and without taking into account either duration dependence or unobserved heterogeneity in the estimation. Therefore, a woman who has her first child in the fifth quarter will contribute 4 observations with dependent variable zero and one with dependent variable 1. The censored observations remaining all 7 years in the sample will therefore contribute 28 quarters with dependent variable zero. This form of estimation allowed us to include time varying covariates that, for example, enabled us to observe whether changes in the woman's labour situation influenced the decision of when to have her first child. This estimation increased the sample to 26490 observations in which each observation is one woman per quarter.

Although most of the women in the sample are active on the labour market or at school, only 39% were employed. Some 19% were unemployed. For working women, we created a dummy taking value 1 if, within the quarter the woman experienced a spell of unemployment, this being a measure of the precarious labour situation many young women experience in Spain. The data show that 11% of employed women experienced spells of unemployment. In the sample a high percentage of women were studying, 35% having not completed their studies and remaining inactive. For this reason, we separated the students from other inactive women, who represent only 7% of the sample.

Different types of workers could be distinguished depending on the type of employment (wage-earners, self-employed and so on) and the type of labour contract. 12% of women were wage-earners with permanent contracts, 11% were wage-earners with fixed term contracts, 1% were wage-earners with other types of working arrangements and 15% were self-employed. It should be noted that the labour composition of the sample can also be related to the sample selection if, for example, women with an adverse economic situation were found to be less likely to have a first child. This type of woman may be overepresented in our sample.

Turning to education, 35% had completed university education, 38% had secondary level and 27% had only finished primary education. The survey provided detailed information on income and wages. In particular, we knew the annual net labour income each adult member of the household received in the previous year and the net household income.

Table 1: Descriptive Statistics.

Variable	**Average (std. error)**
Dependent Variable	
Duration to first child	10.34 (7.17)
Age and Birth Cohort	
Age(1)	25.10 (3.73)
Cohort 1966-69	0.28 (0.45)
Cohort 1970-73	0.38 (0.48)
Cohort 1974-77	0.33 (0.47)
Female Labour Status	
Employed (1)	0.39 (0.49)
Wage-earner with permanent contract (1)	0.12 (0.33)
Wage-earner with fixed-term contract (1)	0.11 (0.31)
Wage-earner with other working arrangement (1)	0.01 (0.11)
Self employed (1)	0.15 (0.36)
Students (1)	0.35 (0.48)
Other inactives (1)	0.07 (0.25)
Unemployed (1)	0.19 (0.39)
Spells of unemployment(1) **(among employed)**	0.11 (0.32)
Completed education	

Primary(1)	0.27 (0.44)
Secondary(1)	0.38 (0.48)
University(1)	0.35 (0.48)
Earnings and income	
Female net hourly wage (1) (2)	4.60 (2.04)
Family Income (3)	4322.75 (3809.28)
Female Wage potential (1) (2)	3.41 (0.85)
Regions (Autonomous Communities)	
Galicia	0.07 (0.25)
Asturias	0.03 (0.17)
Cantabria	0.03 (0.16)
Basque Country	0.06 (0.24)
Navarra	0.03 (0.17)
La Rioja	0.03 (0.17)
Aragon	0.03 (0.18)
Madrid	0.12 (0.32)
Castilla León	0.06 (0.23)
Castilla La Mancha	0.05 (0.21)
Extremadura	0.03 (0.16)
Catalonia	0.10 (0.30)
Valencia	0.07 (0.25)
Balearics	0.02 (0.15)
Andalusia	0.17 (0.38)
Murcia	0.04 (0.19)
Canaries	0.06 (0.24)
# Observations (woman quarter)	26490

Standard Deviation in parentheses. (1) Time varying variables. (2) Wages in Euros per hour. (3) Family income is the quarterly income of the family minus those earnings generated by the work of the woman in the sample.

The wages imputed quarterly could vary from one quarter to the next. As we did not have a similar measure of family income, family income was measured annually. Since our time unit (the minimum period we considered) is the quarter, and employment status information is monthly, we considered an individual as employed in a particular quarter if she worked at least one month in that quarter. The female hourly average wage was 4.66 € whereas the average annual family income was 4197.5 €. Given that our sample is mainly composed of young women, most of whom were living with their parents, the measure of family income recorded a measure of the living conditions of the woman in her parental house, i.e. an Easterlin type variable[3].

Table 2 shows first birth frequencies depending on the labour status of the mother. Although few inactive women in the sample were not students (see Table 1), this was the group of women with the largest exit rate (5.38 % of them had their first child). Employed women returned a rate of 1.8%, whereas unemployed women returned less than half this figure (0.75 %). Students had the smallest rate of exit: just 0.09 % of these women gave birth in the period of observation. Table 2 shows that women working one year before the child's birth accounted for 56% of the births, with inactive women accounting for 29% (the other inactive category excludes students). Another variable of interest is the completed level of education achieved by women. More educated women are more likely to participate in the labour market and will enjoy higher wages. Table 3 shows the exit rates according to the education completed by the mother. Women with primary education had the highest exit rate (1.90%). Women with university education had a higher exit rate than those in secondary education (the highest group), 1.15% as opposed to 0.83%. The increase in education (longer time at school) is one of the reasons given to explain the delay in entering motherhood in some countries (see, for example Gustafsson, Kenjoh and Wetzels, 2001) and may be related to the delay observed in Spain.

Marital status is very important in the decision to have the first child in Spain. In our sample, the exit rate for unmarried women was only 0.6%, as opposed to 4.95% for married women (Table 4). This is consistent with other data in Spain, where only 12.4% of the births in 1994 were to unmarried mothers (FFS data). The decision to get married (or cohabit) and to have a child seem to be closely related and, therefore, studies that analyse the timing of births either control for the endogeneity of marital status (Ahn and Mira, 2001) or restrict the sample to married women (Kalwij, 2000). Heckman and Walker (1991) found that when they controlled for marital status in the fertility equation, the husband’s income became insignificant. This may happen because this

[3] The Easterlin Hypothesis assumes that young people try to maintain at least the living standards they enjoyed in their parents’ house. Numerous cohorts, such as the baby boomers born in the 60s in the USA, will have more difficulties to keep their standards as they will have a more competitive market. So, greater parental income (or lower youth income) will reduce fertility as a reaction to bad life standards.

variable (husband's income) is likely to capture the effect of the omitted marital status variable. If women who have greater preferences for children (or better socioeconomic conditions) have greater probability of marrying, then, including marital status as an explanatory variable of the probability of having a first child will lead to biased estimates due to a problem of endogeneity. For this reason, we include family income,rather than the presence of a partner or the partner's income, as explanatory variables.

Table 2: Quarterly frequencies of first birth according to mothers' labour market status four quarters earlier (% in parentheses)

Woman labour status	Total	Births	Births (%)
Employed	10230 (39.10)	185 (56.57)	1.78
Students	9321 (35.63)	8 (2.45)	0.09
Unemployed	4906 (18.75)	37 (11.31)	0.75
Other Inactives	1706 (6.52)	97 (29.66)	5.38
TOTAL	26163 (100)	327 (100)	1.23

Table 3: Quarterly frequencies of first birth according to completed education of mother (% in parentheses).

Completed Education	Total women	Births	Exit Rate (%)
Primary	7084 (27.08)	137 (41.90)	1.90
Secondary	9908 (37.87)	83 (25.38)	0.83
Tertiary	9171 (35.05)	107 (32.72)	1.15
TOTAL	26613 (100)	327 (100)	1.23

Table 4: Quarterly frequencies of first birth according to marital status (% in parentheses)

Woman marital status	Total	Births	Births (%)
Married	6027 (23.04)	314 (96.02)	4.95
Other marital status	20136 (76.96)	13 (3.98)	0.06
TOTAL	26163 (100)	327 (100)	1.23

4.3 Female wage and opportunity cost of the decision to have first child

Neoclassical models of fertility model the decision-making process of a woman who has to decide about how much of her time she should dedicate to work and how much to childcare (Willis, 1973). These models explain the relationship between fertility and female wages as the aggregate of two opposite effects. On the one hand, a greater female wage imposes a cost on female time: the opportunity cost. This cost is the wage that a woman loses if she decides to dedicate her time to the care of her children or other home tasks, instead of working in the market. On the other hand, an increase in the female wage may have[4] a positive income effect on the probability of having a child, since, if the woman works, she will have more income to dedicate to the care of her children. Most studies find that female wages are negatively related to fertility: the higher the offered wage is, the more women will decide to work and spend less time on childcare. On the other hand, the wage offered in the market cannot be seen as an independent variable of other decisions taken by a woman. Human Capital Theory (Becker, 1964) establishes that wages remunerate productive individual characteristics such as education and years of working experience. Thus, women will decide to increase their human capital (their productive characteristics) if the gains of this increase are greater than the costs. Staying longer in education as a means of improving

[4] According to New Home Economics, higher income may lead to higher or lower fertility depending on the interaction between quantity and quality of children, even if children are a normal good. See Becker, 1981.

future earnings or work stability will also have an effect on other decisions a woman should take. Extra time spent studying will delay the entrance in the labour market and in motherhood. Butz and Ward (1979) found that the relationship between fertility and wages changed from positive to negative according to U.S. data and justified this change by the progressive increase in female participation in the labour market. Ermisch (1989) found that the effect of wages on fertility can be positive when childcare can be purchased in the market. Heckman and Walker (1990) found that the effect of female wages on fertility is negative while the effect of male wages is positive. They also found that the main influence of economic variables on fertility operates through their effect on the decision to have the first birth. Tasiran (1995, 1997) criticizes the use of aggregate wages in Heckman and Walker's study and shows how this relationship is not robust to changes in the wage measure employed. Walker (1997) answered Tasiran by showing that when using microwages, net from measurement error, the Heckman and Walker results stand. Kalwij (1999), using Dutch data in a model with imperfect financial markets, found that couples needed to accumulate savings to be able to afford the fixed costs imposed by children.

Previous studies that estimated the effect of wages on fertility used the observed wages of working women, while not taking account of the wage turned down by non-working women. To take the wages of both types of women (working and non-working women) into account, we estimated a wage equation and imputed a potential wage for each woman to be included in the fertility equation as a means of measuring the effect of wages on the probability of entry into motherhood. This enabled us to measure opportunity cost effects for participants and non–participants.

We applied Heckman's two-step method to estimate the wage equation and the sample we used included women between the ages of 16 and 64 regardless of whether they had a first child or not. In the first stage we estimated a probit where the dependent variable measures the working status of the woman. From this estimation we obtained the Mills ratio that was to be included in the wages equation to correct the selection bias. Heckman (1979) showed that this method provides consistent wage estimates. We show the result of this regression in Table 5. All the coefficients had the expected signs. Age had a positive coefficient but its square had a negative coefficient, indicating that the wage increases with age but at a decreasing rate. The highest wage is achieved at age

46. Our results also show that more educated women obtain higher wages and that married women have higher wages than singles. We used these estimates to calculate the potential wage for each woman according to their personal characteristics. We could not include work-related variables to be able to impute wages to non-workers. Being aware of this limitation, to improve our estimation of potential wages, we used both the average and the variance of the estimated potential wages. To do this, we estimated the variance taking into account the sample selection. How the variance was calculated is shown in Appendix 1. Using both the average and the variance, we generated 100 random distributions and calculated an average of all of them for each woman. The results are shown in Table 6. The average imputed wage for all women was 3.41 € per hour and 3.77 € per hour for workers.

Table 5: Female Wage Estimation

Log (hourly wage)[(a)]	Wage equation	Participation Equation
Age/10	8.96 (25.58)	-1.32 (20.33)
Age^2 /100	-9.57 (-21.75)	
Female completed education (Ref.: primary)		
Secondary	0.29 (18.21)	0.23 (13.26)
Tertiary	0.54 (16.60)	1.07 (60.20)
Demographic Variables		
Number of children (ref: no child)	-0.19 (-3.23)	
Familiy size		-0.10 (-20.64)
Marital Status (Ref.: married)		
Single		-0.09 (-5.28)
Other		0.33 (12.82)
Inverse Mills Ratio	-0.22 (-5.03)	
# Observations	43461	
Censored	29860	
Uncensored	13601	
Wald Statistic	7674.96	

(a) Hourly wage constructed using monthly net wage and number of hours worked.

Regions (Autonomous Communities) and year dummies also included.

Tabla 6: Hourly wage observed and imputed (euros per hour)

Wage	Spain
Workers wage (observed)	4.60 (2.05)

Workers wage (imputed)	3.77 (0.86)
Imputed wage all women	3.41 (0.85)
# Observations	26163
Workers	10716

5. EMPIRICAL RESULTS

Discrete duration model

Using the estimated potential wage from the wage equation as one of the explanatory variables, we estimated the discrete duration model to study the effect of female wages on the decision of entry into motherhood. The dependent variable is the probability of having the first child in each quarter, conditional on not having had one before. The results of our discrete hazard model (equation 8) are shown in Table 7. We estimated the model twice, once using the whole sample of women and a second time restricting the sample to married women. The principal results stand when using this second restricted sample.

In the estimation we controlled for current age and year of birth to account for differences across cohorts. Women who belong to cohorts 1974-77 have the lowest probability of having the first child and women who belong to cohort 1965-69 (the reference group) have the highest. When we control for birth cohort, age reduces the probability of giving birth for the first time.

The decision to have the first child may be related to labour market status and its stability (Ariza, De La Rica and Ugidos, 2005; De La Rica and Iza, 2005). We include labour market status by distinguishing according to the type of contract. Our results show that being a wage-earner with a fixed-term contract and being unemployed significantly reduces the probability of having the first child. We did not find significant differences between inactives and wage-earners with a permanent labor contract. Self-employed women have lower probability of giving birth for the first time than wage-earners with a permanent contract. The variable spells of unemployment are non-significant in any of the estimations when the unemployed variable is included, probably due to some correlation between these two variables.

Education reduces the probability of having a first child, as is clear from the estimated coefficients of the two educational dummies that record having secondary or tertiary education with regard to the omitted category, primary education.

With respect to income and wages, we found that the potential female wage affects the probability of giving birth for the first time positively. Its square affects negatively, showing that a higher wage increases the probability of entering motherhood, although this effect disminishes as wages increase. This result is also obtained when using married women only. The marginal effect of the wage is the proportional change of the hazard rate when the hourly wage increases by 1 € and the rest of the explanatory variables remain at their average values. If we switch from the hourly to the monthly wage for a woman who works 40 hours a week and earns 600 € per month (3.5 €/hour), it means that, for instance, an increase in the monthly wage of 100€ would increase her conditional probability of entering motherhood by approximately 50%. This result suggests us that wage increases lead to positive income effects that dominate negative substitution effects.

Table 7: Discrete Hazard Estimates. Probability of having first child

Variable	All women	Married women
Age and Birth Cohort		
Age(1)	-0.95 (-12.20)	-4.22 (7.09)
Cohort 1970-73	-1.10 (-4.77)	-0.86 (4.76)
Cohort 1974-77	-2.30 (-5.94)	-1.57 (4.71)
Non participants (1)	-0.005 (-0.03)	0.09 (0.57)
Fixed term workers (1)	-0.63 (-2.95)	-0.64 (3.09)
Other workers (1)	-0.37 (-0.82)	0.05 (0.12)
Self employed (1)	-0.43 (-2.05)	0.53 (3.31)
Unemployed(1)	-0.96 (-4.45)	-0.72 (3.64)
Spells of unemployment (among workers)(1)	-0.14 (-0.47)	0.02 (0.06)
Completed education		
Secondary(1)	-7.37 (-16.46)	-2.36 (-6.98)
Tertiary(1)	-17.70 (-17.02)	-5.95 (-8.27)
Earnings and income		
Potential Female Wage (1)	17.87 (13.60)	4.87 (4.76)
Potential Female Wage Square (1)	-0.99 (-8.84)	-0.20 (-2.01)
Family Income(1)	-0.001 (5.52)	-0.0004 (-1.90)
Regional Unemployment Rate(1)	0.16 (8.82)	0.02 (2.15)
Regions (Autonomous Communities)		
Galicia	4.81 (11.09)	
Asturias	-0.75 (1.51)	
Cantabria	2.15 (3.36)	
Basque Country	-0.87 (-2.35)	
Navarra	2.09 (4.77)	
Aragon	2.33 (4.75)	
La Rioja	4.80 (7.57)	
Castilla León	1.09 (3.12)	
Castilla La Mancha	0.70 (2.14)	
Extremadura	1.64 (3.41)	
Catalonia	2.07 (6.68)	
Valencia	3.45 (9.96)	
Balearics	3.28 (7.07)	
Andalusia	-2.63 (-4.55)	
Murcia	4.36 (10.02)	
Canaries	3.47 (9.50)	
Northwest		1.27 (4.17)
Northeast		0.19 (0.68)

Centre		0.77 (3.01)
East		0.72 (2.86)
South		0.50 (1.51)
Canaries		1.35 (4.19)
m2	1.43 (0.41)	
Prob m1	0.18 (0.29)	
Prob m2	0.82 (1.29)	
Log var σ		-3.28 (-0.76)
# Observations (woman quarter)	26490	6341

Z-Statistics in parentheses. (1) Time varying variables. (2) The reference woman was born between 1966 and 1969, is a wage-earner with indefinite contract, with primary education and living in Madrid. (3) Regions in the sample of married women have been grouped: Northwest includes Galicia, Asturias and Cantabria; Northeast includes Basque Country, Navarra, La Rioja and Aragon; Centre includes Castilla La Mancha, Castilla León and Extremadura; South includes Andalusia and Murcia. (4) Unobserved heterogeneity in the estimation with the sample of married women has been estimated using a Gamma function[5] due to convergence problems with the Heckman and Singer type estimates.

Contrariwise, family income affects the probability of giving birth for the first time negatively. This result is not surprising if we take into account that women in our sample are aged 24 on average and that 37% of them are students, living with their parents. Thus, in most cases, rather than measuring a woman's family income, family income actually measures her parents' income. Moreover, students have fewer children. In fact, the negative effect of family income disappears when the sample of married women is used in the estimation.

We also included regional dummies and the regional unemployment rate to obtain a demand side control. We found that women living in regions with higher unemployment rates are more likely to have their first child than women living in regions with lower unemployment rates. This variable may be capturing some other differences across regions in Spain posibly related to a preference for children. We also included regional dummies that were significant in most cases. The reference region is Madrid, and with respect to Madrid, only women living in the Basque Country and Andalusia have a lower probability of having their first child. Women living in Galicia, Cantabria, Navarra, Aragon, Castilla-León, Extremadura, Catalonia, Valencia, Balearics, Murcia and Canaries have a higher probability of entering motherhood.

Bivariate Probit model

Rather than actually measuring the effect of female labour force participation on the probability of giving birth for the first time, we were interested in the effect of wages on this probability. However, given that both decisions—maternity and labour market status—are likely to be taken jointly by a woman, it is important to take into account the potential endogeneity of participation when estimating the probability of having the first child in order to measure the effect of covariates properly. We therefore estimated a bivariate probit where we estimated a woman's fertility and participation decisions together. We defined as participant a woman who was employed or looking for a job (unemployed). The fertility equation refers to the first child only. As in the previous subsection we estimated the model with two different subsamples, one including the whole sample of women (Table 8.1) and a second one restricting the sample to married women (Table 8.2). We also used as a covariate the participation status of the woman one year before (i.e. two years before birth) to measure whether a greater attachment to the labour market influences the maternity decision. We found a positive effect of this variable, although, at 5% level of significance, it was not statistically significant.

As in previous studies on Spain (De la Rica and Iza, 2002 and Adsera, 2004) we found that having a stable job had a positive and significant effect on the probability of having a first child. Relative to age, we found again that women of a given age from later cohorts were less likely to have the first child and, in a given cohort, age also disminished the maternity probability.

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[5] In this case the hazard function is $h_{it} = 1 - \exp\left[-\exp\left\{x_{it}'\beta_t + \gamma_j + \log(\varepsilon_i)\right\}\right]$, where $\gamma_j = \int_{a_{j-1}}^{a_j} \lambda_0(\tau)d\tau$ and ε is Gamma distributed with mean 0 and variance *v*. As the variance of ε goes

Being inactive or unemployed significantly reduces the probability of entering motherhood compared to a permanent contract wage-earner. Having a fixed-term contract also reduces the probability of entering motherhood, as does being self-employed. Spells of unemployment do not have a statistically significant effect. Having completed more than primary education reduces the probability of entering motherhood. This result may be due to the high number of students included in the sample.

The potential female wage increases the probability of participating and once this effect is taken into account, female wages also positively affect the probability of having the first child. Its square affects negatively, as the effect of wage increases, but at a decreasing rate. Although family income has a negative effect on maternity when the whole sample of women is used, it is not statistically significant when only married women are used in the estimation.

zero, the model converges to a proportional hazard with no unobserved heterogeneity.

Table 8.1: Bivariate Probit Estimates. All Women.

Variable	Maternity Equation		Participation Equation	
	Coefficient	Z	Coefficient	Z
Age and Birth Cohort				
Age(1)	-0.14	(-7.73)	0.03	(2.38)
Cohort 1970-73	-0.30	(-5.24)	0.07	(2.16)
Cohort 1974-77	-0.49	(-4.84)	-0.24	(-4.69)
Employment Status				
Participation	0.07	(1.39)		
Non participants (1)	-1.41	(-15.43)		
Fixed term workers	-0.20	(-2.96)		
Other workers (1)	-0.13	(-0.89)		
Self employed (1)	-0.11	(-1.65)		
Unemployed(1)	-0.22	(-3.32)		
Spells of unemployment	-0.02	(-0.20)		
Completed education				
Secondary(1)	-0.55	(-5.42)	-1.09	(-16.90)
Tertiary(1)	-1.68	(-8.06)	-1.56	(-10.56)
Earnings and income				
Potential Female Wage(1)	1.55	(5.09)	1.64	(8.47)
Potential Female Wage Square (1)	-0.09	(-2.86)	-0.09	(-5.13)
Family Income(1)	-0.0002	(-3.54)	-0.0003	(16.53)
Regional Unemployment			0.02	(6.43)
Partner (1)			0.25	(10.05)
Correlation	-0.81	(-11.85)		
# Observations	26490			

Standar Deviation in parentheses. (1) Time varying variables. (2) Year and region dummies also included as controls.

Table 8.2: Bivariate Probit Estimates. Married Women.

Variable	Maternity Equation		Participation Equation	
	Coefficient	**Z**	**Coefficient**	**Z**
Age and Birth Cohort				
Age(1)	-0.11	(-4.64)	-0.07	(-3.12)
Cohort 1970-73	-0.24	(-3.67)	0.03	(0.48)
Cohort 1974-77	-0.32	(-2.63)	-0.13	(-1.21)
Employment Status				
Participation	0.04	(0.90)		
Non participants (1)	-1.38	(-18.48)		
Fixed term workers	-0.18	(-3.08)		
Other workers (1)	-0.04	(-0.29)		
Self employed (1)	-0.14	(-2.17)		
Unemployed(1)	-0.18	(-2.66)		
Spells of unemployment	-0.001	(-0.01)		
Completed education				
Secondary(1)	-0.74	(-6.32)	-0.32	(-2.54)
Tertiary(1)	-1.65	(-6.63)	-1.01	(-3.73)
Earnings and income				
Potential Female Wage(1)	1.37	(3.76)	1.29	(3.27)
Potential Female Wage Square (1)	-0.08	(-2.27)	-0.05	(-1.22)
Family Income(1)	-0.0002	(-0.20)	-0.0003	(0.06)
Regional Unemployment			0.03	(6.24)
Correlation	-0.93	(-7.50)		
# Observations	6341			

Standar Deviation in parentheses. (1) Time varying variables. (2) Year and region dummies also included as controls.

Finally, the rho coefficient that estimates the correlation between the two equations is negative, as was to be expected, showing that fertility and participation are negatively related. Otherwise, when participation and fertility are estimated simultaneously and wages of working and nonworking women are taken into account, we did not find the negative effect of wages on fertility found in many other papers (Butz an Ward, 1979; Heckman and Walker, 1990; Walker, 1997, among others). This result suggests that the positive income effect overcomes the negative substitution effect. For a country like

Spain, where female employment rates are still below other European countries and where the unemployment rate is too high, we think that this result may shed some light on the delayed maternity decision for Spanish women. Our results suggest that, at least in Spain, a country with high unemployment and where young people and especially women experience major difficulties in finding a stable job, higher female wages could increase the probability of having the first child, the positive income effect of wages dominating the negative substitution effect.

6. CONCLUSIONS

We studied the probability of giving birth for the first time for a representative sample of young Spanish women. We focused our attention on the effect of female wages on the probability of having the first child. The wage is the opportunity cost of spending an hour taking care of children instead of working. Previous studies using the observed wages of working women only found that the female wage had a negative effect on fertility. Given that non-working women also have a non-observed offered wage that they have turned down, we believe it is important to take into account the wages of these non-working women when estimating the effect of wages on fertility. Therefore, we estimated a potential wage for each woman in the sample, independently of her labour market status, and used this potential wage to study the effect of female wages on the probability of having a first child. This potential wage was calculated controlling for sample selection. Further, given that job characteristics cannot be included in the wage equation to improve our estimation of potential wages, we used both the average and the variance of the estimated potential wages to calculate the imputed wage.

In line with previous articles on Spain, we found that unemployment or having temporary contracts had a negative effect on the probability of giving birth for the first time. We also found, contrary to previous studies, that in Spain the *female potential wage* affects the probability of entering motherhood positively. This result suggests that the income effect of a wage increase overcomes the substitution effect.

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APPENDIX 1: MEAN AND VARIANCE OF WAGES

In our estimation of potential wages, we used the standard Heckman's two-step method to control for the sample selection problem of observing only the wage of those currently working and not observing the wage offered to those who have turned down the offer. The two equations of the model are the following wage and participation equations:

$$w_i = x_i^{'}\beta + u_i \qquad (1)$$

$$I_i = z_i^{'}\gamma + \varepsilon_i \qquad (2)$$

The main problem of using this method to impute a wage to all women is that we do not observe labour variables for non-workers and then we have to exclude this variable from the regression. However, we used both the mean and the variance of the Heckman estimates to obtain a potential wage as accurate as possible, instead of the standard use of the mean on its own. Since we only observe the wage conditioned on being in work, the mean wage is (Heckman, 1979):

$$E(w_i / x_i, z_i, I_i > 0) = x_i^{'}\beta + E(u_i / x_i, z_i, I_i > 0) \qquad (3)$$

Taking advantage of the fact that the correlation between u_i and ε_i is not zero, $u_i = \sigma_{u\varepsilon}\varepsilon_i + \zeta_i$, we have: $E(w_i / x_i, z_i, I_i > 0) = \sigma_{u\varepsilon} E(\varepsilon_i / x_i, z_i, I_i > 0) = \sigma_{u\varepsilon}\lambda(z_i^{'}\gamma)$ (4)

then:

$$E(w_i / x_i, z_i, I_i > 0) = x_i^{'}\beta + \sigma_{u\varepsilon}\lambda(z_i^{'}\gamma) \qquad (5)$$

where $\lambda(z_i^{'}\gamma)$ is known as the **inverse of Mill's ratio**.

Then, taking into account the selection, the variance is:

$$Var(w_i / x_i, z_i, I_i > 0) = \sigma_{u\varepsilon}^2 \left[1 - z_i^{'}\gamma\lambda(z_i^{'}\gamma) - \lambda^2(z_i^{'}\gamma)\right] + Var(\zeta_i) \quad (6)$$

where the term in brackets can be estimated using:

$$z_i = \alpha + \beta \, z_i^{'} \lambda(z_i^{'} \gamma) - \delta \lambda^2 (z_i^{'} \gamma) + e_i \qquad (7)$$

where $z_i = \left(w_i - \hat{E}(w_i / x_i, z_i, I_i > 0\right)^2 \qquad (8)$

With the values of the correct mean and variance of wages for all women, we generated its distribution for every woman, assuming normality. From these distributions we made a random extraction of 100 values and an average of all of them per woman. This is the measure that we used as the imputed wage in our study.