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Research Article

**Women's wages and childbearing decisions:
Evidence from Italy**

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Women's wages and childbearing decisions: Evidence from Italy

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Abstract

During the early 1990s, Italy became one of the first countries to reach lowest-low fertility. This was also a period in which women's education and labour force participation increased. We analyze the role of women's (potential) wages on their fertility decisions by making use of two different surveys. This enables us to apply discrete-time duration models. For first births, we find evidence of non-proportional hazards and of some "re-cuperation" effects; for second and third births, instead, wage exhibits small intensity although there is a clear division between Northern and Southern Italian regions.

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1. Introduction

During the early 1990s, Italy became one of the first countries to reach lowest-low fertility, i.e. a Total Fertility Rate (TFR) below 1.3. Whereas the reasons behind the Italian fertility decline are many, a prominent explanation suggests that it has been driven by increasing opportunity costs. Certainly, the strong fertility decline during the nineties was followed by a significant increase of education and labour force participation among women. Given this setting it is of interest to establish the extent to which women's wages are linked to their fertility decisions. As a result we analyze fertility behavior of Italian women during the period 1983 to 2003. We pay particular attention to the role of potential incomes (or predicted wages), which in turn depend directly on women's educational attainment, and therefore their opportunity cost, and we investigate how this affects their childbearing decisions.

So far, research on this issue in Italy has been hampered by lack of data sources that contain information on both fertility behavior and women's income or earnings. We overcome this problem by combining two different data sets. Data from the Italian Institute of Statistics Labour Force Survey of 2003 (ISTAT-LFS hereafter) are used to reconstruct women's fertility histories. Potential wages (i.e. hourly labour income) are derived from the income and earnings data from the Bank of Italy's Survey on Household Income and Wealth (BOI-SHIW). Data from the BOI-SHIW are used to predict women's potential wage, which is then introduced as an explanatory variable in discrete time hazard regression models for first, second, and third births.

Our analysis includes women aged between 15 and 40 years in 2003. Their predicted wages depend on a range of background variables, including their educational attainment. As such the predicted wages reflect their potential wage, which in turn is a measure of their socio-economic characteristics. Our interest lies in assessing the extent to which socio-economic characteristics, expressed through predicted wages, may play a key role in the postponement of motherhood in Italy and their impact in the transition to first, second and third birth. In particular, we assess whether potential wages are the only determinants in these transitions, or whether other socio-cultural factors may be in play for delayed motherhood and second and third birth decisions.

The remainder of this paper is organized as follows: in Section 2 we review the existing literature and describe the main features of the Italian setting as far as childbearing decisions are concerned. Section 3 provides a description of the data and methods we use. In Section 4 we present our main results, summarized in Section 5.

2. Wage and childbearing decisions

2.1 Previous literature

It is useful to start from an economic perspective when considering the role of wages and income on fertility. Becker (1960) argues that there should be a positive correlation between the number of children and the household income. Highly educated mothers, however, tend to substitute child quality for number of children (see Becker and Lewis (1973)). Since both production and rearing of children are time intensive, an increase in wage rates induces a negative substitution effect on the demand for children (see for instance Becker (1965); Mincer (1963)). To this extent, theoretical research on fertility (such as Hotz, Klerman, and Willis (1997); Becker (1981); Willis (1973)) shows that women's income is negatively associated with childbearing as a higher income implies higher opportunity cost of children. In other words, with high earnings, it becomes more expensive for her to take time away from work to rear children. In general, the opposing income effect is unlikely to outweigh the negative substitution effect. Overall, therefore, we would expect the effect from women's wages to be negative. For men, in contrast, the income effect tends to dominate since they spend less time on rearing children, though the magnitude of these effects will vary across countries and birth parity (Butz and Ward 1979; Willis 1973).

Ermisch (1989) provides an important extension to the Beckerian argument, the insight being that women's fertility decision does not only depend on their wages, but also on availability of external child-care. In the presence of costly external child-care, women with very high earnings, traditionally having a high positive opportunity cost, may instead have more children, because they are more able to afford external child-care. Those with very low wages are less likely to afford external child-care, but may still have higher fertility due to low opportunity costs. In this scenario both low and high income women will have more children, whereas those with middle income will have lower fertility (i.e. lower demand for children). The argument depends, of course, on the availability of child-care. Italy is characterized by a lack of affordable child-care when compared to the Scandinavian countries. Moreover, we might expect that the effects of child-care might be lesser in Scandinavia than in Italy, as the Scandinavian countries (within each country) have far more uniformity in the availability cost of child-care, due to the government's role in establishing and pricing such facilities. Because of the nonhomogeneous institutional background that depicts a Northern Italy with a daycare system stronger than that of the South - a South with working women mostly involved in public jobs and a North with a much higher prevalence of private firms (Del Boca, Locatelli, and Vuri 2005) - regional differences in Italy play an important role in explaining fertility decisions.

In recent years research has shifted towards investigating the timing of births rather than completed fertility (see for instance Heckman and Walker (1990a)). In these studies, hazard rate models are used to empirically analyze the timing and spacing of births. Most of the studies show that, compared to women with low opportunity costs, women with high wages (i.e. high opportunity cost of having children) have births later. However, a source of potential conflicting patterns is present among high-earning postponers. Higher incomes can mean both higher opportunity costs and a better ability to pay for high cost private child-care, as previously discussed. There is also ample evidence to suggest that the presence of children has a significant negative impact on the woman employment probability (see Mroz (1987); Heckman and Macurdy (1980)). Women with high earnings stay longer in education, and possibly in the labour market, before they have the first child. For this reason, return to work might be a more affordable option for high-earning women, while others are more likely to give up. As Pronzato (2009) shows this can be due to the lack of protected leave, low childcare availability and attitudes on gender roles and family-work orientation.

A prominent part of the existing literature argues that women's responsibility for child-rearing may reduce her time in paid work. Joshi (1990), for instance, analyzes how work patterns can be different, in terms of switching from full time to part time work or not employed, by comparing mothers with childless women. The causal direction is explained through the impact of children on the women's job market opportunities and their level of income. Miller (2010) and Waldfogel (1998) offer similar perspectives.

Likewise, the decision to enter parenthood has important time dimensions to it. A couple may want to pay attention to the cost of having children *early*, comparing it with the cost of having them *later*. In such a way, the optimal age of having the first birth can be seen as a trade-off between investment in human capital and career planning (see Gustafsson (2001)). As emphasized by Cigno and Ermisch (1989), Cigno (1991) and Gustafsson and Wetzels (2000), it is important to take into account possible consequences to lifetime earnings given different scenarios of birth timing and spacing. In light of this it may be optimal for many women to delay motherhood until the opportunity cost of child-care (with respect to her career) has decreased, which implies completion of education and getting a foothold in the labour market, before entering motherhood.

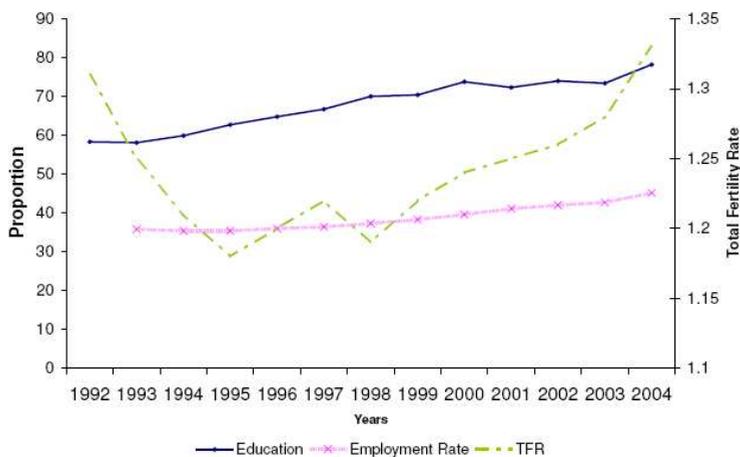
Formally, the timing of first birth is a function of the opportunity cost of time and the foregone human capital cost (Gustafsson 2001; Cigno 1991). However, the effect of income on the timing of births may work through different paths. Gustafsson (2005) suggests that, for Swedish young individuals, the postponement of formal education works by delaying couple formation rather than by delaying parenthood once the couple is formed. So, family policies may have a pro-natalist effect in allowing Swedish couples to have the first child earlier. A somewhat different interpretation is offered by Amuedo-Dorantes and Kimmel (2005) asserting that college-educated mothers can profit from delaying mother-

hood, because they are in a position to negotiate a family-friendly work environment with flexible work schedules. Similar evidence is shown for the British case (Kneale and Joshi 2008).

2.2 The Italian setting

Figure 1 provides stylized facts of fertility in Italy. Two prevalent trends are depicted for the last two decades in Italy: a decline in total fertility and a steady increase in women's educational attainment, together with higher female employment rates (see also Figure 2). Using a sensitivity analysis, Rindfuss, Guzzo, and Morgan (2003) showed that a 1% increase in female labour force participation was associated with a 3.15% decline in Total Fertility Rate. During the same period, the mean age of onset of motherhood increased from 25 years in 1980 to almost 27 in 1990, reaching a level of 28 years in 1995 and 28.7 in 1997 (Figure 2).

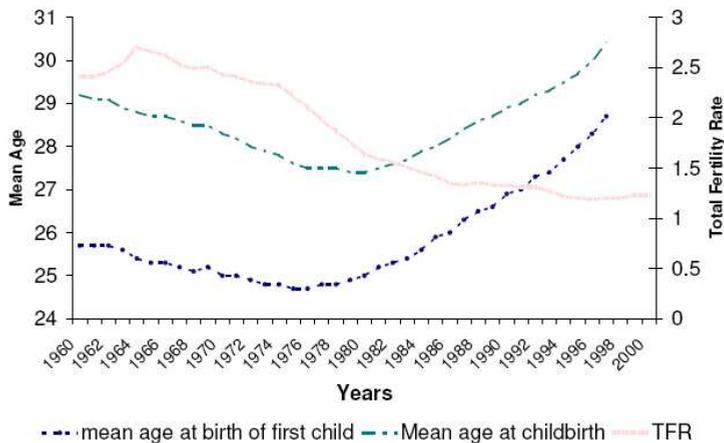
Figure 1: Education, TFR and female participation rates in Italy (1992-2004)



Notes: Fertility rates: 1998 and 1999 are provisional values and 2000, 2001, 2002 are estimated values. Education is the percentage of the female population aged 20 to 24 having completed at least upper secondary education. The female employment rate is calculated by dividing the number of women aged 15 to 64 in employment by the total female population of the same age group.

Source: Eurostat.

Figure 2: Mean age at first birth, at childbirth and TFR in Italy (1960-2000)



Notes: Mean age of women at childbearing is the mean age of women when their children are born. For a given calendar year, the mean age of women at childbearing can be calculated using the fertility rates by age (in general, the reproductive period is between 15 and 49 years of age).

Source: Council of Europe (2001).

As Kohler, Billari, and Ortega (2002) point out, Italy was, together with Spain, the first country to reach the threshold of so-called lowest-low fertility, i.e., below 1.3 children per woman. It is well established that the emergence of the lowest-low fertility in Southern Europe is not connected to any steep increase in childlessness (Billari and Kohler 2004). Available parity-specific data on fertility show that most of the fertility decline in Italy during the last twenty years is due to a decreasing progression to the second, third and subsequent children. As a consequence, the probability of having a first child has not changed in spite of the tremendous economic and social changes of the Italian country during that period (see Dalla Zuanna 2004). Moreover, the personal ideal family size for around 60% of Italian women aged 20-34 years is two children; while one quarter has a preference for large families (Goldstein, Lutz, and Testa 2003).

Economic factors might thus shape a) the timing of first birth more than the probability of ever having a first birth and b) the timing and the probability of ever having births of higher order. Data from opinion surveys support this idea. In a recent survey (2002) on a sample of mothers aged around 42, living in five Italian cities, women gave reasons they had stopped at the parity they actually experienced. Concerning the non transition to the

third birth (i.e., women who stopped at two children), economic reasons were cited as important for women who experienced a worsening of their financial situation after the birth of the first and second child. Many women argued that monetary transfers for the first three years after the birth of a third child, or a lower but longer financial incentive, could have changed their decision to stop at parity two (De Santis and Breschi 2003). Although being possibly biased as an ex-post motivation, this role of economic factors is specific for third birth.

3. Data and methods

Micro evidence on the relationship between the timing and the spacing of births has been scarce for the Italian case, the main reason being a lack of data sources provided with wage and income information linked with women's fertility histories, where the latter has sufficient sample size. The data requirement for estimating the impact of wage on the timing of births is demanding: we need panel data with a long time dimension or retrospective data on the complete employment and fertility histories. Usually, and this is our case, a researcher has available only cross section data or panel data with a short time dimension. Fertility histories may be reconstructed on the basis of the age of the children in the household but the complete labour market history of the women in the household will be more difficult or even impossible to reconstruct. This is because one observes, for each household, the number of children and the employment status of the women in the household only at the time of the interview.

Unfortunately none of the currently available Italian data sets contain all the required information. The BOI-SHIW provides detailed information on employment and income of family members, labour market activities, payment instruments and forms of savings, socio-demographic characteristics of the household. However, the sample size is too small to conduct fertility analysis, particularly for third births. The ISTAT-LFS, 2003, contains detailed information on the family structure, labour market, work experience, part time and full time employment. The main drawback of this survey is that it does not collect information on household earnings and income. The sample size is however large and suitable for fertility analysis.

In order to overcome these limitations, we combine the two data sets. We use the BOI-SHIW data set to estimate earnings equations, from which the predicted earnings are matched to women in the ISTAT-LFS data set.

3.1 The Survey of Households' Income and Wealth (BOI-SHIW)

The BOI-SHIW started in 1964 and it is provided with an annual and historical version. Because a long time span is analyzed, this paper uses the historical version of the survey which is the main source of information on Italian household savings and wealth. The survey collects information on more than 22,000 individuals per year (around 8,000 households) with 13,500 individuals receiving an income. From the survey we have information on payroll employees, members of professions, family business and pensioners. Respondents report their annual wage (both if they are part and full time workers), the number of months and hours worked on average per week, so that we can construct an hourly wage measure independent from the opportunity cost of the woman dividing the annual net wages by the annual working hours.

The survey is provided with net wages, excluding tax and social security contribution, but including month salary ("13 month" and "14 month" salary), bonuses or special allowances (family allowances, productivity bonuses, sales commissions, etc.).

Our unit of analysis is the woman. We only consider households in which it is possible to link every woman with their co-residing children. In total this produces a sample of 20,003 individuals in 2002 with 4,749 women in their childbearing stage. In other words, we consider women aged less than 40 years, who are not pensioners and not studying.

3.2 The Italian Labour Force Survey (ISTAT-LFS)

The ISTAT-LFS, implemented since 1959, has become a continuous survey in 2005. It was a quarterly survey up to 2004; for each year four waves are carried out. The survey collects information on more than 300,000 households, representing around 800,000 individuals (1.4% of the total national population) distributed over 1,351 municipalities (out of 8,100). It assesses the state of the Italian labour market, supply side only. The sample design is a two stage rotating sample design with stratification of the primary units (municipalities). Each household is included first in two waves, then left out for two waves and then included in another two waves. The ISTAT-LFS offers different sections dealing with demographic characteristics of the households, present job (with all the information taken from the month before the interview), job experience, looking for a job, and relationship with public employment centers. However, there are no retrospective fertility histories available. Instead we know the number and ages of children living in the households in which the woman is either the household head or the spouse of the household head.

Our sample considers all the households included in the four waves carried out in 2003, in which the most recent information is taken, if the household is studied for two consecutive waves. Fertility histories are derived from recorded co-residing children at

the time of interview. Thus, we are able to reconstruct retrospective fertility histories for each woman back to 1983. We make the assumption the woman had a “regular” schooling career, as the dataset is provided to with the maximum level of education and the year this was achieved. It is reasonable assume that a woman drops out when she spends more than a “regular time” in an educational level. We also assume that the women in the dataset had no geographical mobility, a plausible assumption as the women’s internal migration was almost stable in the years 1988-2003 and attained 2% on average per year (Istat, General Population Register).

The sample includes a small number of adopted children or stepchildren, but we exclude any offspring who might have died or moved away. In Italy, mortality at adult age is low, children of divorced parents are almost exclusively living with their mothers, and a very low proportion of young individuals leave the parental household before the age of 23 (Aassve et al. 2002). Nevertheless, in order to ensure that the recorded children are the only ones of the mother, we limit the analysis to only include women who are aged 40 or less in 2003. Given that the mean age of leaving home is rather high in Italy (Aassve et al. 2002) this selection does not reduce the sample by much. We also exclude households in which we were unable to link children with mothers (i.e. male head of the household with no wife and all single men). In this way we end up with a sample of 34,129 women. In total, the fraction of women that can be matched with their co-residing children is 95%. Unfortunately, we are not able to use information of husbands since we only know the marital status of women at the time of interview. This problem also applies to widows, in that we do not know when the husband died.

3.3 Descriptive statistics and methods

Our methodological strategy can be summarized in three steps. First, using the historical version of the BOI-SHIW data, we estimate separate wage equations for the years 1991, 1993, 1995, 1998, 2000, 2002, assuming that for the period 1983-1991 wages evolved as they did in 1991. For the missing years, mean wages are interpolated from the adjacent years. Second, we match predicted wages onto the ISTAT-LFS data using common characteristics of women in both data sets. Third, we create three sub-samples consisting of women being at risk of the first, second and third birth.

The details of estimation of wage equations is given in the appendix. The discussion of the relevant ISTAT-LFS descriptive statistics, using wages predicted from the BOI-SHIW, is reported in Table 1. We are assuming that a woman who is relatively poor in her 20s need not to be so in her 40s.

Given our sample of women aged 15 to 40, Table 1 shows that the mean age at first birth is 26.2 years for those who have had a first birth already, when this mean is evaluated over the sample period from 1983 to 2003. Because of the nonhomogeneous institutional

background, the regions are stratified into North, Centre and South. While poor women in the North (where poor means an hourly wage lower than the 25th percentile of the wage distribution for the whole country) start to have children early, women in the South tend to delay motherhood until they are 23.5. Transition to second (and respectively third) birth takes place on average two years after the first (second) child is born. Low (high) wage mothers living in the North tend to concentrate their fertility history in a shorter (longer) time span: they start at 23.0 (31.2) and end it after 4 (2) years, conditional on reaching the threshold of the 3rd birth.

Table 1: Mean age at first, second and third birth across regions and wage

	All	Wage \leq 25th percentile	Wage \geq 75th percentile
<i>Transition to 1st birth:</i>			
	26.2	23.4	31.3
South	25.0	23.5	31.5
Centre	26.7	23.3	31.9
North	27.1	23.0	31.2
<i>Transition to 2nd birth:</i>			
	28.9	26.4	31.7
South	27.9	26.5	31.5
Centre	29.5	25.6	32.7
North	29.9	25.6	31.5
<i>Transition to 3rd birth:</i>			
	30.8	29.1	32.9
South	30.4	29.2	33.6
Centre	31.3	27.8	33.4
North	31.4	26.7	32.7

Notes: The percentiles refer to the distribution of predicted hourly wage from the estimated wage equation as reported in the appendix. The means are for closed birth interval only (computed only for women experiencing a birth); women are classified according to the wage rate they had at birth. The 20 Italian regions are classified according to the following: Piemonte, Valle d'Aosta, Lombardia, Liguria, Trentino, Veneto, Friuli Venezia Giulia, Emilia Romagna belong to the North. Toscana, Umbria, Marche, Lazio are considered Central Regions. Abruzzi, Molise, Campania, Puglia, Basilicata, Calabria, Sicilia, Sardegna belong to the South.

Source: Own calculation from ISTAT-LFS, 2003, with fertility histories retrospectively reconstructed from 2003 back to 1983.

The estimation of the impact of predicted and time-varying wage $\hat{\omega}_{it}$ on fertility is implemented through a set of discrete time event models. Consider a series of P predictors $X_{1ij}, X_{2ij}, \dots, X_{Pij}$ and let x_{pij} denote individual i 's values for the p -th predictor in time j . The hazard function is defined as:

$$h(t_{ij}) = Pr[T_i = j | T_i \geq j \text{ and } X_{1ij} = x_{1ij}, X_{2ij} = x_{2ij}, \dots, X_{Pij} = x_{Pij}] \quad (1)$$

that is, the population value of discrete-time hazard for person i in time period j is the probability that he/she will experience the target event in that time period, *conditional* on no prior event occurrence *and* his or her particular values for the P predictors in that time period (Jenkins 1995).

We estimate a variety of single spell discrete time duration models (Singer and Willett 2003) where the baseline hazard for the transition to first birth is a function of the age of the woman while the one for the transition to second (third) birth is a function of the spell, duration since first (second) birth. In order to assess the impact of predicted wage on fertility, different specifications of the model are offered: we start from a simple one controlling for wage and its square and extend it to allow for some interaction between covariates, capturing both duration and cultural/institutional effects.

The omission of relevant information for the husband could cause a serious bias in our model, as fertility decisions are expected to depend on the income level of the household. It has been proved that the omission of unobserved heterogeneity in single spell discrete time duration models causes a rescaling of all coefficients (Nicoletti and Rondinelli 2010). It has also been shown that the rescaling factor is close to one when using time varying coefficients. Given that wage is a time varying regressor the omission of the unobserved heterogeneity do not cause any relevant bias in the estimated parameter.

In all our discrete time duration models (estimated from ISTAT-LFS), wages are predicted from BOI-SHIW. The standard errors are consequently corrected for the fact that we use generated regressors in the fertility equation (see Pagan (1984); Arellano and Meghir (1992)). To do so we use the bootstrap method with 300 replications and observations clustered at the woman level. Any further increase in replications does not significantly affect the estimated standard errors.

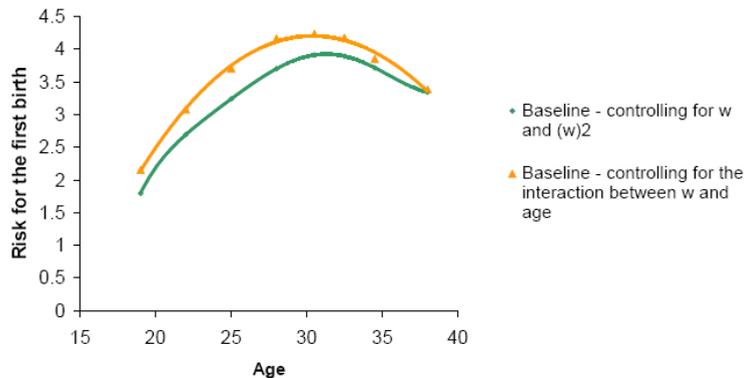
4. Empirical results

In this section we show how predicted wages may affect the risk of having the first, second and third birth. The estimations ignore the effect of men's wage, since we only know the marital status of the woman at the moment of the interview and detailed information on partner's income is not included in the data.

4.1 First birth

As expected, the baseline hazard for first births depicts an inverse U-shape. The effect of wages is approximated by a second order polynomial, which simply means that we include wages and its squared term in the regression. Column (1) of Table 2 shows that the general effect of women's wage is negative: the higher the wage, the lower is the risk of entering motherhood (-0.344). The effect of wage is, however, nonlinear which is confirmed by the positive coefficient associated with the square term. Controlling for wage and its squared term only, the maximum of the parabola ranges between 30 and 31 as depicted in Figure 3, meaning that the mean age at first birth is between 30 and 31 years of age. This effect is, however, overestimated. In 1997, for example, the average age at first birth was 28.7 (see Council of Europe (2001)). As shown in Figure 3, once we control for interactions between age and wage, the baseline hazard is shifted to the left (an anticipation effect), resulting in a correct mean age for the first birth being between 27 to 28.

Figure 3: Baseline for the first birth



Notes: Women between 15 and 18 are the reference group.

Table 2: Transition to first birth and wage effect

Variable	Coeff. (1)	(Std. Err.) (2)	Coeff. (3)	(Std. Err.) (4)	Coeff. (5)	(Std. Err.) (6)
<i>General wage effect:</i>						
Wage	-0.344**	(0.006)				
(Wage) ²	0.066**	(0.003)				
<i>Wage and age:</i>						
Wage × age (15-17)			-0.759**	(0.060)	-0.688**	(0.062)
Wage × age (18-20)			-0.540**	(0.019)	-0.475**	(0.019)
Wage × age (21-23)			-0.557**	(0.014)	-0.492**	(0.015)
Wage × age (24-26)			-0.434**	(0.012)	-0.371**	(0.015)
Wage × age (27-29)			-0.292**	(0.012)	-0.236**	(0.016)
Wage × age (30-31)			-0.072**	(0.019)	-0.023	(0.021)
Wage × age (32-33)			-0.054*	(0.024)	-0.004	(0.027)
Wage × age (34-35)			0.078*	(0.037)	0.125**	(0.040)
Wage × age (36-40)			0.123*	(0.051)	0.172**	(0.052)
<i>Region effect:</i>						
Centre					-0.206**	(0.023)
North					-0.242**	(0.020)
<i>Wage and region:</i>						
Wage × Centre					-0.040*	(0.019)
Wage × North					0.008	(0.016)
Constant	-5.735**	(0.052)	-6.132**	(0.108)	-5.935**	(0.112)
No. of observations	425,014		425,014		425,014	
No. of women	34,129		34,129		34,129	
No. of spells	25,233		25,233		25,233	
Pseudo R^2	0.086		0.090		0.091	
Log-L	-87524.64		-87159.352		-87064.898	

Notes: Discrete-time logit hazard regressions. Baseline (i.e. age coefficient) not reported. Reference category for region is South. Bootstrapped standard errors in parentheses. Standard errors are computed using 300 bootstrap samples.

† $p < 0.1$; * $p < 0.05$; ** $p < 0.01$.

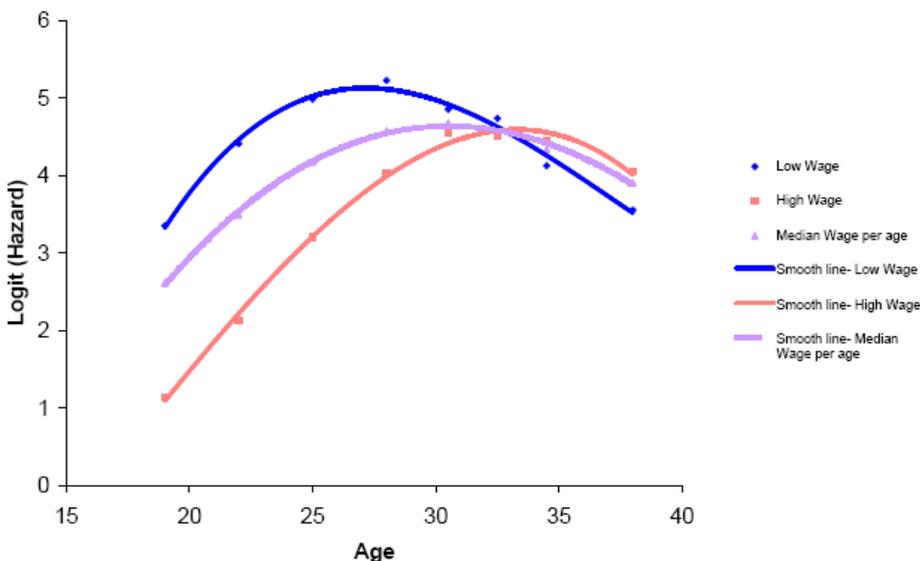
In columns (3) and (5) of Table 2, we report the estimated coefficients for women being at risk of the first birth where wage has been stratified by age and interacted with region. This allows us to understand how wage impacts on first birth after controlling for age and cultural or institutional effects. Column (3) of Table 2 shows that wage has a strong negative effect on the risk of the first birth for young women, the effect becoming closer to zero at the age of 34-35. Women with high wages have a higher likelihood of postponing motherhood which is confirmed by the importance of wages in driving the onset of motherhood. When controlling for regions (column (5) of Table 2), we mostly find the same negative pattern for younger women but the turning point is now between 32-33 years. After 34 years, increasing wage by one unit increases the likelihood to experience motherhood. Women in Central and Northern Italy are less at risk of first birth when compared with women in the South. However, when controlling for cultural and institutional effects (Wage and region effect of Table 2), high wage Central women are less likely to become mothers for the first time compared to the Southern one.

In order to better understand how different wage levels affect the timing of first birth, we simulate the hazard paths for a low and high wage woman. We choose two extreme situations. A low (high) wage woman is assigned a wage set equal to the 10th (90th) percentile of the hourly wage distribution. Both paths are plotted together with the hazard of the median wage woman in Figure 4. The Figure shows a non proportional hazard: low wage women are at a higher risk of experiencing the first birth when they are very young, reaching the maximum level when they are aged between 25 and 30 years of age. High wage women tend to delay, maximizing the likelihood for the first birth when they are 30 or older. The median wage woman has, not unexpectedly, an intermediate position. When the three paths cross (at age 32) low, median and high wage women experience similar risks though with the difference that a low wage woman has already reached her maximum, a median woman is at the maximum risk level and the high wage woman is yet to reach it.

We also plot the simulated survival curves in Figure 5 by: a) identifying a group of women with low (high) wage defined as those women with predicted wage less than the 25th percentile of the wage distribution for that age (greater than 75th percentile of the wage distribution); b) calculating the predicted survival curves based on the coefficients estimated in column (3) of Table 2. Given that wage is a time dependent covariate the survival curves are plotted only for women who have been low (high) income earners throughout their whole lives. A highly interesting pattern here is the strong recuperation effect among high wage women. Though the recuperation is not complete, the Figure suggests that high wage women are in effect quite similar to low wage women in terms of completed fertility of first parity. In other words, high wage women start the onset of motherhood later, but recuperate after some time. This pattern is rather stable, and persists also when high wage women are defined according to the 90th percentile as opposed to

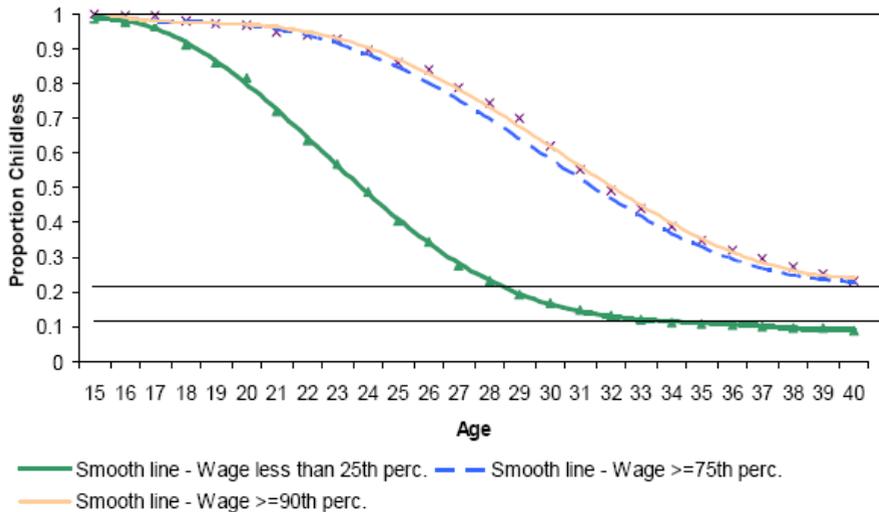
the 75th percentile. The same is true if low wage women are defined according to the 10th percentile. We find that higher wages have their primary effect on postponement of first birth so that when wages are higher, pregnancy tends to be concentrated in a shorter span of the life cycle. This is in line with the opportunity cost theory (see Heckman and Walker (1990a)): few employed mothers would quit their job to have children if an exit from the labour market could seriously damage their future labour market prospects (Boix 1997). This effect would be more important the greater the uncertainty in the labour market or the higher the unemployment rate. This is one of the reasons why Central and Northern regions exhibit negative estimated coefficients for the risk of the first birth when compared to the South, where there is higher labour market uncertainty. In the Italian setting, characterized by an imperfect labour market (protecting those who are in, penalizing those who are out), starting late or discouraging a career can be very costly in terms of earnings so that it makes a lot of difference whether the woman is or is not employed, or a least has, or has never, been employed.

Figure 4: Timing of first birth for a low and high wage woman



Notes: Logit (Hazard) estimated from Table 2, column (3). Low and high wage are constant through age: low (high) wage are set equal to the 10th (90th) percentile of the hourly wage distribution.

Figure 5: Predicted survival curves. Postponement and recuperation effect



Notes: Median predicted survival per ages across wages (25th, 75th, 90th percentile). The proportion of women childless is estimated from Table 2, column (3). Recuperation effect is the vertical distance between the predicted survival curves.

The paths may also be explained in terms of recruitment policies which differ across regions. As many firms often recruit young talents prior to motherhood rather than women in the post-parental phase, this might reinforce postponement incentives (Gustafsson 2001). The negative coefficients of the interaction between wage and region (column (5) of Table 2) can therefore be explained by the large presence of public jobs in the South where working women are mostly employed in teaching or administrative activities, mainly driven by a lack of private firms. Women working in public sectors tend to have permanent contracts and pre- and post-natal paid periods, providing greater stability (i.e. low likelihood experiencing job redundancy). The North, in contrast, is characterized by a much higher prevalence of private firms. To this extent women from the North, normally with higher wage levels, are induced to postpone motherhood until they reach a position in their career planning. Consequently, depreciation of women's human capital due to temporary absence from the labour force due to childbearing, might be more severe compared to women in the South (Bernardi (2003) offers a qualitative research on the role that social influence plays on the transition to motherhood in Italy).

Another argument which is consistent with the negative interaction between wage and region is that, presumably, mothers with higher career prospects decide to postpone maternity until they obtain a more stable labour market situation (e.g. getting a permanent contract, see De La Rica and Iza (2005) for an explanation for the Spanish case). Unfortunately, our data reconstruction of women's job history makes it difficult to verify this hypothesis. However, it seems there is no institutional or cultural effect driving the first birth decision. If this was the case, we should have found a positive coefficient for the interaction of wage with North where a more efficient child care system is set up in order to help working mothers to reconcile job and family decisions.

4.2 Second and third birth

The fertility hazard is modeled in such a way that a woman becomes at risk of a second birth once she experienced the first; similarly, she becomes at risk of the third once she has experienced the second. A trivial consequence of this obvious fact (yet important for completed fertility) is that the age at first birth determines the age of becoming at risk of a second birth. However a delay in the first birth does not necessarily imply lower completed fertility as a woman may have an incentive to accelerate the second and third birth to achieve her desired fertility level. Our estimates in Table 3 prove that the higher the age of first birth, the less is the risk of second birth (-0.017).

The coefficient for the age of second birth is smaller than the one for first birth, meaning that the risk of having a third child is strongly negatively correlated with the age at which second birth took place. In other words, mothers who tend to delay first birth are at a lower risk of having a second birth and those who delay the second have lower likelihood of having a third. Wage has a negative effect on fertility at a decreasing rate. The main difference is that while the wage effect for second birth is negative and significant, the coefficient for the third birth is near zero and insignificant, indicating that the wage effect plays a marginal role in third births (Table 3). A merely tentative explanation of the weaker effects of wage on second and third birth is that some women may in fact have no wage, having left the labour market after the first birth, while others may have lower wages, if they extended their stay at home over the third month, and others may have their full wage. Unfortunately, no information on the parity specific employment pattern of mothers is available which would help to buttress this finding.

Table 3: General wage effect for second and third birth

Variable	Coefficient (Standard Error)		Coefficient (Standard Error)	
	Transition to second birth		Transition to third birth	
<i>General wage effect:</i>				
Wage	-0.071**	(0.009)	-0.03	(0.018)
(Wage) ²	0.052**	(0.004)	0.098**	(0.011)
Age at first birth	-0.017**	(0.003)		
Age at second birth			-0.095**	(0.006)
Constant	-2.786**	(0.078)	-1.910**	(1.900)
No. of observations	122,608		83,949	
No. of women	25,227		14,483	
No. of spells	14,309		2,503	
Pseudo R^2	0.440		0.024	
Log- L	-42221.856		-10989.240	

Notes: Discrete-time logit hazard regressions for second and third birth. Baseline not reported. The number of women at risk of a second birth differs from the number of spells reported in Table 2 because women having twins as first births are not in the risk set of second birth as they are at risk of a third birth. Bootstrapped standard errors in parentheses. Standard errors are computed using 300 bootstrap samples.

† $p < 0.1$; * $p < 0.05$; ** $p < 0.01$.

Looking more specifically at how wage may affect transition to second and third births, as reported in Table 4, we find wage has a negative impact on the probability of second birth one to two years after the first birth (column (1) of Table 4). This is either because households prefer to use additional resources on the first child or because it is the moment women tend re-enter the labour market. Even if their wage increases, they have little incentive to use it for rearing another child. The situation is different after three to four years. In this case an increase in wage has a positive effect on second birth. A similar pattern is followed by women at risk of third birth: the wage effect is negative immediately after second birth and close to zero three to four years after the second birth. This suggests that the timing for second and third births is positively induced by socioeconomic variables only three to four years after having experienced previous birth. Afterwards, the risk of giving a new birth decreases as we move away from the date of the preceding birth. This is perfectly in line with Heckman and Walker (1990b) point of view concerning 3rd birth in Sweden: when wages are higher there is a primary effect in reducing third birth and a secondary effect that forces pregnancy to be concentrated in a shorter span of life cycle.

Table 4: Wage and region effect for the transition to second and third birth

Variable	Coefficient (Standard Error)		Coefficient (Standard Error)	
	(1)	(2)	(3)	(4)
	Transition to second birth		Transition to third birth	
Age at first birth	-0.026**	(0.003)		
Age at second birth			-0.101**	(0.006)
<i>Wage and duration:</i>				
Wage × dur (1/2)	-0.050 [†]	(0.029)	-0.399	(5.588)
Wage × dur (3/4)	0.088**	(0.016)	-0.090	(0.073)
Wage × dur (5/6)	-0.049**	(0.018)	0.058	(0.048)
Wage × dur (7/9)	-0.135**	(0.024)	-0.069 [†]	(0.042)
Wage × dur (10/13)	-0.124**	(0.041)		
Wage × dur (10/14)			-0.166**	(0.045)
Wage × dur (14/25)	0.103	(0.079)		
Wage × dur (15/24)			-0.360**	(0.085)
<i>Region effects:</i>				
Centre	-0.455**	(0.030)	-0.270**	(0.084)
North	-0.444**	(0.025)	-0.119 [†]	(0.066)
<i>Wage and region:</i>				
Wage × Centre	0.003	(0.030)	0.166*	(0.080)
Wage × North	0.116**	(0.021)	0.181**	(0.054)
Constant	-2.258**	(0.080)	-1.679**	(13.730)
No. of observations	122,608		83,949	
No. of women	25,227		14,483	
No. of spells	14,309		2,503	
Pseudo R^2	0.050		0.024	
Log-L	-41986.374		-10987.727	

Notes: Discrete-time logit hazard regressions for the transition to second and third birth. Baseline not reported. Reference category for Region is South. The number of women at risk of third birth differs from the number of spells reported for the transition to second birth because women having twins as second births are not in the risk set of third birth. Bootstrapped standard errors in parentheses. Standard errors are computed using 300 bootstrap samples.

[†] $p < 0.1$; * $p < 0.05$; ** $p < 0.01$.

Table 4 shows that women living in the Centre and North are less likely to experience a second or third birth compared to the South. But Table 4 also reveals a clear institutional or cultural effect underlined by the interaction of wage and region. An additional unit of wage for Northern women increases her risk to have a second or third birth. One might argue that this reflects confidence in the child-care system among working mothers living in the North and Centre (see Del Boca, Locatelli, and Vuri (2005) for a complete explanation). In Southern Italy, crèches are not widespread and working mothers tend to prefer informal child care, i.e. baby-sitters and grandmothers living nearby. When any of these two conditions is not satisfied, having another child means additional costs in terms of searching for childminders. The decision to have a second or third birth, even if in presence of higher wage, is therefore not straightforward (see also Diprete et al. (2003)). The situation is different in the North, where the large presence of public and private services offers a more reliable child-care system that also facilitates child-care even when children are very young (Del Boca, Locatelli, and Vuri 2005). We could also think there might be a selection effect at work. Only a few high wage women in the North have two children: if they are selected to for high fertility preferences, they tend to have a third child (see Kravdal (2001) for a different setting).

5. Summary and concluding remarks

The aim of this paper is to find an empirical connection between women's potential wages and both the delay in motherhood and the transition to second and third birth in Italy.

Using two different data sets, we estimate the effect of predicted wages on the postponement of motherhood. First, a wage equation including detailed controls for women's educational attainment is estimated using a Tobit model. Second, we use the predicted wage as a regressor in discrete time hazard models for the transition to first, second and third births. We find that the wage effect is negatively correlated with having children. The magnitude, however, varies according to the birth order. Wage has a strong negative effect in the timing of first birth. Consistent with opportunity cost theories, women with higher wage tend to delay motherhood. Our estimates suggest a non-proportional hazard and a strong, though incomplete, recuperation effect. Women with higher wages start having children later, but recuperate after some time, while poor women have children earlier. Furthermore, there is no evidence that institutional or cultural effects are responsible for the postponement of maternity.

The pattern is different for second and third births. The wage effect is negative, but it has smaller intensity compared to the first birth. The coefficient for third birth is close to zero, confirming that wage plays a relatively small role. Nevertheless, we find evidence of a clear North/South division in the decision to have more than one child. We might argue

that this reflects the availability of a good childcare system among Northern and Central working mothers, as described in the empirical analysis of Del Boca, Locatelli, and Vuri (2005) for the Italian setting. As a possible explanation, an additional unit of wage for a Northern woman could increase the risk of experiencing a second and third birth because of the childcare provision.

Our results provide interesting evidence of the key role that women's wages play in fertility decisions.

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Appendix

Wage equations are commonly estimated (see for instance Mincer (1974)) by linear regressions where the dependent variable is the natural log of the reported wage. Typical variables to be included are age and age squared, education (either in term of years of education or as a dummy variable reflecting the educational level), type of education, work experience, number of children, age of the children, ethnicity, region, and profession. In theory, we can estimate separate wage equations for men and women within the status of working persons and pensioners. We do exclude pensioners: women of pensionable age are excluded as we limited our analysis to those who are 40 or less. Male are not considered because of the established unit of analysis. Husbands' wage is ignored since we only know the marital status in 2003 and information about the wedding date is not provided in the data. However, there will be women who are recorded with zero wage simply because they do not work. The problem is that they might have chosen not to work because they would receive relatively low wage. In terms of the economic theory, they do not work because their offered wage is lower than their reservation wage (i.e. the lowest wage for which they would chose to work). If we choose those who work in our wage equation only, we do get a selection bias. One standard solution for this problem is to estimate a participation equation using a probit model (the so-called generalized Tobit or Tobit of second type, see Gourieroux 2000).

Consider the market wage (Y_{mi}) and the reservation wage (Y_{ri}) of person i specified as following:

$$Y_{mi} = x_i'\beta + \epsilon_i \quad (2)$$

and

$$Y_{ri} = z_i'\gamma + \nu_i. \quad (3)$$

For the market wage equation (2), x_i' indicates the set of covariates, β the corresponding parameters, and ϵ_i the error terms. For the reservation wage equation (3), z_i' is the set of covariates, γ the corresponding parameters, and ν_i the error terms.

The observed wage is given by:

$$Y_i = \begin{cases} Y_{mi} & \text{if } Y_{mi} \geq Y_{ri} \\ 0 & \text{Otherwise} \end{cases}$$

with:

$$Y_i = \begin{cases} x_i'\beta + \epsilon_i & \text{if } \epsilon_i - \nu_i \geq z_i'\gamma - x_i'\beta \\ 0 & \text{Otherwise.} \end{cases} \quad (4)$$

The two step estimation provides a first step for the estimation of a probit model for the probability to be in the labour market and a second step for the estimation of

the regression model with an additional variable (the inverse Mill's ratio) using the sub-sample of individual with $d_i = 1$, where:

$$d_i = \begin{cases} 1 & \text{if } Y_{mi} \geq Y_{ri}, \quad \text{i.e. a woman works} \\ 0 & \text{if } Y_{mi} < Y_{ri}, \quad \text{i.e. a woman does not work} \end{cases}$$

In our case, we are prevented from using this standard solution because of the combination of two different and independent data sets. Mapping the predicted BOI-SHIW wage into the ISTAT-LFS data set requires having exactly the same variables in both data sets. Including information about the husbands as regressors (i.e. z_i/t) in the participation equation attenuate the bias of the coefficients, but we impute, to the ISTAT-LFS women, a predicted wage which captures, year by year, the presence of an additional wage. As we only know the marital status of the woman at the time of the interview, the husband's wage (i.e., an additional wage) is ignored.

Given this setting, the lack of data available and the data mapping, we estimate a Tobit model (a Tobit model of the first type or Tobit with deterministic censure, see Tobin (1958)) which censors the wage distribution at a selected (upper or lower) point. The woman wage equation is left censored at zero.

Define:

$$\begin{cases} Y : & \text{Hourly Observed Wage} \Rightarrow \min(Y)=0; \\ a : & \text{Positive Constant}; \\ W = Y + a : & a - \text{augmented Hourly Wage} \Rightarrow \min(W)=a; \\ w = \lg W : & \Rightarrow \min(w)=\lg a. \end{cases}$$

Given this definition,

$$Y_i = \begin{cases} Y_{mi} & \text{if } Y_{mi} \geq 0 \\ 0 & \text{if } Y_{mi} < 0. \end{cases} \tag{5}$$

we are interested in taking \lg of (5). Taking a positive constant a and assuming that $W_i = Y_i + a$, we get:

$$W_i = \begin{cases} a + x_i/t\beta + \epsilon_i & \text{if } x_i/t\beta \geq -\epsilon_i \\ a & \text{Otherwise} \end{cases}$$

with $\min(W_i) = a$ and:

$$w_i = \lg(W_i) = \begin{cases} \lg(a + x_i/t\beta + \epsilon_i) & \text{if } x_i/t\beta \geq -\epsilon_i \\ \lg(a) & \text{Otherwise} \end{cases}$$

and $\min(w_i) = \lg a$.

$\lg a$ is the lower bound for the log hourly wage distribution and it is the point of censure of the Tobit model. All the observations with augmented log hourly wage less than $\lg a$ are censored. If we apply to this model a standard OLS, the estimates are biased and inconsistent (Tobin 1958).

Table 5 presents the estimated coefficients from an hourly wage equation for women in a Tobit model using the BOI-SHIW data set in 2002.⁴ The selected sample in Table 5 contains 4,749 women in the condition of being mothers, 2,639 of these women are censored because they exhibit a zero wage.

As we expected, the coefficient for age is positive and for age squared is negative: this is in line with the economic and econometric literature that wage increase with age but at decreasing rate (Mincer 1974). Our results confirms the classical way to take education as a proxy of wage (Schultz 1985). Moving from lower levels of education to higher ones induces a gain in terms of hourly wage rates. We do not include the position and the activity status of a woman in the because it is not provided in the ISTAT-LFS. To be more precise, the ISTAT-LFS asks women position only at the time of the interview and we are no able to reconstruct it back in 1983. The hourly wage does not exhibit very large variability across different region of Italy, even if there is a clear tendency of lower hourly wage in the South (linked both with the presence of public job and the percentage of women not working) when compared to wages in the North.

⁴Wage equations estimated for other years are not reported and are available upon request.

Table 5: Estimated hourly woman wage equation (Tobit model)

Variable	Coefficient	(Standard Error)
<i>Age:</i>		
Age	0.308**	(0.022)
(Age) ²	-0.004**	(0.000)
<i>Education:</i>		
Middle School	0.459**	(0.096)
High School	1.421**	(0.096)
Bachelor's Degree	2.264**	(0.120)
Post-Graduate Qualification	3.916**	(0.749)
<i>Regions:</i>		
Piemonte	1.377**	(0.160)
Valle d'Aosta	1.144	(0.712)
Lombardia	1.248**	(0.141)
Trentino	1.764**	(0.236)
Veneto	1.133**	(0.157)
Friuli	1.254**	(0.233)
Liguria	1.342**	(0.188)
Emilia Romagna	1.603**	(0.158)
Toscana	1.107**	(0.166)
Umbria	1.276**	(0.246)
Marche	1.211**	(0.222)
Lazio	1.011**	(0.153)
Abruzzi	0.551*	(0.250)
Molise	0.525	(0.344)
Campania	-0.475**	(0.175)
Basilicata	1.264**	(0.248)
Calabria	0.635**	(0.219)
Sicilia	0.016	(0.168)
Sardegna	0.543*	(0.224)
Constant	-7.322**	(0.517)
No. of women	4,749	
No. of censored women	2,639	
Pseudo R^2	0.14	
Log- L	-5929.3336	

Notes: Dependent variable is the woman log hourly wage. Equation estimated from the BOI-SHIW, 2002.

Reference group for education: none and elementary school. Reference group for regions: Puglia.

† $p < 0.1$; * $p < 0.05$; ** $p < 0.01$.

