Differences in Delaying Motherhood across European Countries:

Empirical Evidence from the ECHP

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Abstract

In this paper, we aim to assess some of the potential determinants of first childbirth timing in Europe, in a comparative perspective, using micro-data from the European Community Household Panel Survey. We follow the demographic approach to decompose the differences between rates, in the part due to the national population composition by specific characteristics, and in the part due to different propensities for women with given characteristics. Specifically, we show what the probability of entering into motherhood would be for Italian women if they had the same human capital, the same level of labor participation, the same timing in terms of education and first job start as women living in another European country. On the basis of our results, we discuss the possible effect on comparative fertility of policies favoring changes in the work and educational characteristics of Italian women to allow them to be more like women in other European countries.

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1. Introduction

In the last decades a general and progressive delay of the first childbirth has been observed virtually in every European Union country. The percentage of births to mothers aged thirty or over exceeds 40% in various countries, including Sweden, Denmark, Norway, Finland, Netherlands, Italy and Spain (Pinnelli and De Rose 2001). This is considered one of the most characteristic feature of fertility change in Europe, so that some authors refer to it as a distinctive "postponement transition toward a late-childbearing regime" (Kohler et al. 2002).

As supported by micro level analysis (Morgan and Rindfuss, 1999; Billari and Kohler 2000), the delay of the first child birth is a relevant cause of reduction of completed fertility. Undeniably, the compression of the reproductive span may affect the possibility for women to fulfil their desired level of fertility, due to probable sub-fecundity or even sterility impediments (Ongaro 2003).

Since the timing of the first childbirth seems so important to explain subsequent fertility behaviour, we study its determinants in a comparative perspective across 10 European Union (EU) countries, using the European Community Household Panel Survey (ECHP). We focus on the impact on entry into motherhood of variables such as educational level, age when the highest level of education was completed, and age at first job. Specifically we estimate, separately for each country, a discrete time hazard model with time varying covariates (see Allison 1982, Yamaguchi 1991, Petersen 1991, Petersen and Koput 1992, Jenkins 1995 and Sueyoshi 1995). This allows us to consider to some extent the dynamics between education, labour participation and childbearing decisions. The dynamic approach is of paramount importance for evaluating the time when a birth minimizes - or at least reduces – costs arising from women's forgone earnings and depreciation of human capital during maternity.

According to the literature - based on income maximisation framework (Gustafsson 2001, Hotz *et al* 1997, Happel *et al.* 1984) – women with a higher degree of human capital and a shorter work experience are more likely to delay motherhood or to remain childless. However, recent micro-level studies have shown contradictory empirical evidence. For instance, highly educated or career women seem to enter motherhood earlier in the Northern European countries (Kravdal 1994, Hoem 2000, Andersson 2000). Conceivably, these ambiguous findings might reflect substantial cross-country differences, which we would like to point out.

Therefore we conduct a decomposition analysis of the differences in the first birth hazard rates across Europe in two components (Kitagawa 1955). The first component is due to the population composition by specific characteristics, whereas the second one is due to different specific rates for women with given characteristics. Our decomposition analysis differs from the conventional demographic approaches because we do not use sample relative frequencies to estimate the specific hazard rates, rather we consider the parametric discrete-time duration models estimated to analyse the determinants of first childbirth timing. This allows us to avoid the curse of dimensionality (or sparseness of the data) problem arising when using a large set of covariates. Moreover, thanks to this dynamic modelling approach we are able to consider time-varying variables such as spacing between leaving school, starting of first job and first childbirth timing.

Specifically, we focus our attention on Italy, where the postponement of the first childbirth is particularly evident, and we apply a decomposition analysis to explain the differences in the harzard rates between Italy and any other of the nine European countries considered in the analysis. We wonder what would be the probability of entering into motherhood of the Italian women if they had the same human capital, the same level of labour participation, the same timing in terms of education and first job start as women living in another country. Theoretically, we try to answer to the following question: can policies favouring a change of the work and educational characteristics of the Italian women toward other European countries help in reducing the gap with respect to the rest of Europe? We specifically focus on the so called "tempo policies", i.e. those policies that are aimed at affecting the tempo fertility, through for instance efforts in shortening the average school duration and shifting the timing of education at younger ages (Lutz and Skirbekk 2004).

The answer to the above question is not straightforward. This is because considering an Italian woman with education and work experience characteristics closer to those of another European country does not ensure that her childbearing propensity will become similar, too. In other words, the gap between the hazard rates for the timing at first childbirth observed for two different countries can be due to differences both in the observed characteristics and in the coefficients of the estimated hazard models. Differences in the coefficients reflect an intrinsic difference in the childbearing propensity, which persists even if everything else - at least everything we control for - is equal. Those differences may be due to unobserved heterogeneity across countries and in particular to different cultural and policy contexts. If the differences in the hazards are mainly due to differences in the variables, tempo policies can plausibly have an impact on the probability of being mothers. If the gaps between hazards are instead mainly due to the coefficients, then their effect is more ambiguous.

The rest of the paper is organized as follows. Section 1 reviews the theoretical approaches explaining motherhood postponement. Section 2 describes the data and the variables used in the study. Section 3 discusses the statistical model used to analyse the determinants of the timing at first birth and presents the estimation results. Section 4 introduces the decomposition analysis and discusses the results of its application to explain the differences in the timing at first birth between Italy and any other country. Finally, Section 5 draws some conclusions.

2. Theoretical background

Many studies have been dedicated to assess the determinants of motherhood postponement. Most of them have shown that deferral of motherhood is the last consequence of a more general delay in almost any step in the so-called transition into adulthood (Ongaro 2003), including the timing of sexual initiation, leaving parental home, and entering a union. Livi Bacci (2001) sees the signs of a "delay syndrome" in all this, particularly in Southern European countries.

Several tentative explanations have been offered with regard to postponement. Some of them stress the relevance of economic and structural constraints (e.g. Happel *et al*. 1984), some others emphasise the socialisation process, the changes occurred within the family and a social values shift (e.g. Schizzerotto and Lucchini 2002, Lestaeghe 1995). We focus exclusively on the economic approach, which assumes that reproductive behaviour is the outcome of a rational choice process and individuals have almost complete control over fertility. It is then sensible to hypothesise that the choice of having a child at a specific time requires an evaluation of costs and benefits related to motherhood in a long, as well as in a short, term perspective. Therefore, the likely future economic situation and expected income profiles should be taken into account in this evaluation.

Postponing motherhood, to when there are fewer uncertainties about the economic situation and union stability, allows one to evaluate more precisely costs and benefits of childbearing (Kohler *et al.* 2002, Simò *et al.* 2002). However, delaying is not a cost-free decision: as the wished age at motherhood increases, women approach their biological limit. This increases demands on medical assistance – e.g. in vitro fertilization - and raises biomedical expenses (Wetzels, 1999, ch. 7). In addition, late mothers are subject to more substantial risks for their own health and for their late born child (Gustafsson, 2001). It is conceivable that the evaluation of latter type of costs is less relevant, as they are related to a more remote future, but it is also evident that they may even offset the gains in lifetime earnings, especially if the cost to pay is an involuntary permanent childlessness. The question then arises: how can one minimise the risks and costs related to maternity? The economic theory provides an answer to the question of the optimal age at motherhood basically through two main explanations.

The first one - known as the consumption-smoothing motive (Hotz et al. 1997, Happel et al. 1984) - emphasises the role of man as main earner in the household and the preference for a smooth consumption over the life cycle. This theoretical explanation suggests that the best time to become parents is when the household income is the highest. As the male partner is considered the breadwinner, household income is likely to be the highest when male partner income is (Happel et al 1984). If men earnings increase over time, then the life cycle utility is maximized by delaying childbirths to the female biological limit. In conclusion, when women's maternity leave and child related expenses are delayed to a time when partners' earnings are relatively high, the household smoothes its consumption profile and increases its lifetime utility.

The second explanation is known as the career planning rationale. If a woman wishes to pursue a labour market career, she will have to complete education and find a stable and adequate employment. Both education and work experience can be considered as an investment in human capital, as they improve job carrier opportunities. Unavoidably, childbearing compels women to temporarily withdraw from the labour market. This determines a short-term loss of resources, due to income loss, but it also has consequences for women’s human capital and future earnings profile. Some author refers to that as the “childbearing penalty” (Joshi 2002), which is a component of child cost strongly linked to mother’s age. Assuming that, during job interruptions, women’s human capital depreciates, the longer the job interruption, the greater the probability of human capital devaluation, and the higher the wage penalty. To be more specific, the human capital loss depends on the pre-maternity human capital level, on the rate of depreciation due to non-use and on the life income profile (Gustafsson 2001). The steeper the wage increase (linked to the work experience) is, the longer the postponement of motherhood should be (Wetzels 1999, Cigno and Ermish 1989). In other words, women with a flat wage profile by duration of work experience have less incentive to postpone motherhood, with respect to the ones who expect to have high wage increases. Therefore, postponement can be seen as a rational response to socio-economic constraints. However, Happel *et al.* (1984) find contradicting results: they observe delayed first births when wages are rather insensitive to work interruptions.

We follow suggestions from the career planning approach to specify our model for the timing of the first birth. In particular, we consider the effects of the variables related to women’s career, namely work experience duration, and human capital, namely number of years spent in education, on the timing of first birth.

3. Data and variables

The European Community Household Panel Survey (ECHP) is a rich longitudinal micro data source providing comparative socio, demographic and economic variables at household and individual level for all countries in the European Union. We use the ECHP user data base, called shortly ECHP-UDB, released in 2003 and covering the period from 1994 to 2001. We consider only 10 countries, namely Belgium, Denmark, France, Germany, Greece, Ireland, Italy, Portugal, Spain and the UK. For comparability reason we keep only countries, which participated to the ECHP for all 8 waves, so that Austria and Finland are excluded from the analysis. The Netherlands is not considered, because some of the variables used in our analysis are not available for this country. Finally, we drop Luxembourg because of its small sample size.

We select a sub-sample of 23,221 women born between 1954 and 1977 (17 to 40-year-old at the first interview), regardless of their marital status and parity. At each wave, women in the ECHP are asked to answer some retrospective questions, including some on the age at which they completed their highest level of education, the age at which they had the first work experience and the age at which they gave birth to their first child[1]. For all women with a child in 1994 (the first wave of the ECHP), we consider the above retrospective information collected in the first wave. Women without a child in 1994 are instead followed until 2001, the last wave of the panel or until they give birth. However, the lack of very detailed retrospective data on fertility and job career does not allow a complete reconstruction of life event histories.

Information in the database allows us to consider the following time-varying variables:

1. level of education measured by the age when the highest level of education was reached.[2]
2. age,
3. time elapsed since the highest level of education was completed,

[1] The age at which women give birth to their first child is identified by using dates of birth of women and of children living together with women. For this reason we exclude from the sample the women with children living away in the first wave of the panel. Anyway, the percentage of women excluded for this reason is low, about 1% of the sample.

[2] This choice is determined by comparability reasons, as many differences in the educational systems across countries make the education levels – as classified in the ECHP –not always strictly comparable.

4. work experience measured as the time elapsed since starting the first job experience,
5. a dummy indicating if a woman has ever worked,
6. a dummy for women still in education.

The choice to focus only on past experience, instead of considering also current activity status, is dictated by data availability. We cannot observe the labour work history for women who were older than 17 in the first wave of the panel. Anyway, focusing on the subsample of women having their fist child during the panel and considering them only for the time interval 1994-2001 covered by the ECHP, we find that current activity status and income have an insignificant impact on the hazard of having a first birth after controlling for the level of education, timing of leaving education and start of first job.

Again for data availability reasons, we do not consider the women's marital and cohabiting statuses. Focusing again on the above subsample of women, we find that living in a couple is an important determinant of childbearing but, for some countries, as for example Spain, out-of-wedlock childbearing occurs so rarely that it is impossible to identify the effect of being single on first childbirth hazard probability. If we assume that living in a couple is a prerequisite to first childbirth for women older than 17, then neglecting the dummy indicating whether a woman lives in a couple or is single is justified.

In the literature, women's education level – a proxy of human capital and of earnings potential - has usually been found to play a pivotal role in determining the timing of first births, as well as lifetime fertility. It has been generally found that higher education delays motherhood. What it is not completely clear is whether there is a sort of simple mechanical effect on delay, linked to an increase in the number of years spent in education, and/or a real inhibiting effect on motherhood due to a different propensity of more educated women.

High education levels have a pure "mechanical effect" on maternity postponement when the delay is just equal to the number of additional years spent in education compared with lower education levels. However, the delay may be longer because highly educated women:

1. have usually a stronger job attachment (Bratti 2001), a stronger preference for career over maternity (Gustafsson 2001) or are self-selected among women being less family-oriented;
2. are likely to invest more time in job search, after finishing school, in order to find a more satisfactory job (Gustafsson, Kenjoh and Wetzels 2001);
3. are likely to pay a higher cost if they decide to have a child at the beginning of their career (Wetzels 2001).

Gustafsson and Wetzels (2000), Ermish and Ogawa, (1994) and Rindfuss *et al.* (1988) find a large effect of education level on timing of the first birth. Kravdal (1994) seems instead to find little effect of education level on entry into motherhood once controlled for union status, age and work experience. Therefore, the impact of education level appears to be mainly due to shorter work experience observed for higher educated women. A similar result has been obtained in Gustafsson, Kenjoh and Wetzels (2001). They find that the effect of education disappears once they measure the timing at first birth as the duration since finishing education. For the above reasons, we consider also time elapsed since the highest level of education was completed in order to assess whether a prolonged education has a pure mechanical effect completing studies at different ages.

Work experience is included in our models, using - as a proxy - the time elapsed since the entry into the labour market. It is necessary to point out that – as stated before - we are not able to reconstruct all the details of women's job career and therefore we do not know the number of job interruptions and their length. Nevertheless, we consider the duration since the beginning of the first job important because the entry into the labour market is a relevant step in the path towards economic independence, autonomy and adulthood. Both theoretical and empirical studies have produced contradictory predictions and estimates of women's work experience effect on entry into motherhood. Cigno and Ermish (1998) state that a longer work experience tends to accelerate first birth, even if their predictions do not seem to be well-supported by British data. Happel *et al.* (1994) predict instead an inhibiting effect of work experience in the USA. Blossfeld and Huinink's findings (1991) support the latter hypothesis in Germany, while Kravdal (1994) finds that the first birth risks increase sharply after the fourth year of working experience, but after six years a plateau is observed for Norway.

One of the clearest relations – observed in the literature – is that being still in education inhibits entry into motherhood (Andersson 2000, Hoem 2000 and Beets et al. 2001). Students' lower first birth rate may be a consequence of too low income to afford childrearing costs. Moreover, student mothers cannot rely on appropriate policies (e.g. family allowances), which are usually destined to employed women (Andersson 2000). It is also possible that students perceive that childbearing inhibits transition to higher educational levels (Hoem and Hoem 1987), which in turn has a negative effect on women's human capital and lifetime earnings capacity. In addition - more simply - student lifestyle may not fit with family responsibility (Gustafsson *et al.* 2001).

In Tables 1 and 2 simple descriptive statistics are reported. It is immediately evident that Italy represents a peculiar case: the percentage of women still childless, the mean age at first birth, as the age at first job, is the highest compared to other countries, while on average the age of completion of education is not among the highest. Furthermore, over a third of the sample never entered the job market in Italy, while in other countries these proportions are lower, and in a few cases less than 10%.

3. Age at first birth: statistical model and results

In this section, we describe briefly the statistical model used to analyse the determinants of the age at first birth and we present the results of the model estimation for Belgium, Denmark, France, Germany, Greece, Ireland, Italy, Portugal, Spain and the UK using the ECHP.

3.1 The statistical model

The statistical model used to explain the age at first motherhood is a discrete-time duration model. In other words, we estimate the probability of transition to motherhood at a specific age, say t, for childless women at age ($t-1$), where age is measured in years.[3] Henceforth we will use indifferently transition, hazard or first birth probability (or rate) to refer to the probability of transition to motherhood at a specific age.

For all women with a child in 1994, the first wave of the ECHP, we identify the age at their first child birth by considering the date of birth of their oldest child. Women without a child in 1994 are instead followed until 2001, the last wave of the panel. For women still childless at last wave or before dropping out of the panel we consider the problem of the right censure assuming that the censure is not informative.[4] If a woman, childless in the first wave,

[3] For more details on discrete-time duration models we refer to Allison (1982), Yamaguchi (1991), Petersen 1991, Petersen and Koput 1992, Jenkins (1995) and Sueyoshi (1995). We do not consider a continuous time survival model because we measure the age of women at first childbirth in years. This is because it is possible to measure the age in months only for some countries included in the ECHP.

[4] We say the right censure is not informative if the duration of the woman's participation in the panel and the duration until her first childbirth are independent of each other, everything else given. See Lancaster (1990) for more details.

drops out from the panel but is responding in last wave, we update the information using the last wave questionnaire to identify the potential childbirth occurred in the meanwhile.

We explain the probability to give birth to a child for a woman at the age t given that she was childless at the age $t-1$ using her personal retrospective information on age at the completion of the highest level of education, *age at hle*, and age at the first job. More precisely, using those two variables we build a set of explanatory variables for each woman from the age 17 to the first childbirth or to the right censure. Each woman has a number of observations equal to the number of years between the age of 17 and the first childbirth or the right censure. The set of explanatory variables consist of:

1. age and age square (*age*^{2}): t and t^2;
2. age at which the highest level of education was completed (as a proxy of education level), say *Age at hle*, which takes value *0* if women are still in education;
3. number of years since completion of the highest level of education, say *Age-Age at hle*, and (*Age-Age at hle*)^2, which take value *0* if women are still in education;
4. dummy variable taking value *1* if a woman is still in education and *0* otherwise, (*Still in education*);
5. number of years since the beginning first work experience, *Age-Age at 1st job* and (*Age-Age at 1st job*)^2, which take value *0* if women have never worked;
6. dummy variable taking value *1* if a woman has never worked and is not still in education and *0* otherwise, say *Never worked.*

Finally, we build the dependent variable, say $r_{i,t}$, for the generic woman i-th at age t, which is a dummy variable taking value *0* for each year the woman is childless from the age of *17* onward and *1* when the woman age is equal to her age at first childbirth.

We estimate a discrete-time hazard model for the duration from age 17 to the first childbirth. The estimation of a discrete-time hazard model consists in the estimation of a sequential binary model for the dummy $r_{i,t}$. We consider a probit model,[5] so that the hazard function can be written as

$$\Pr(r_{it} = 1 | r_{it-1} = 0, X_{i,t}, t) = \Phi(\beta X_{i,t} + \alpha_t),$$

where $X_{i,t}$ are the above described variables for the i-th woman at age t, β is a vector of parameters of interest, α_t is an age-specific intercept. In the empirical application we assume that α_t is a quadratic polynomial function in the age, i.e. $\alpha_t = a + b\, t + c\, t^2$.

Assuming that $(r_{i,t}|r_{i,t-1}=0, X_{i,t})$ be identically and independently distributed (i.i.d.) across women, the parameter of interest β can be estimated by maximising the product of the likelihoods for each woman in the sample. For a woman giving birth to her first child at the age t, the likelihood is:

$$\Phi(\beta X_{i,t} + \alpha_t) \prod_{s=17}^{t-1} \left(1 - \Phi(\beta X_{i,s} + \alpha_s)\right).$$

While for a woman childless in the last wave of the panel, say at the age t, the likelihood is:

$$\prod_{s=17}^{t} \left(1 - \Phi(\beta X_{i,s} + \alpha_s)\right).$$

A similar likelihood is also valid for a woman who drops out of the panel before giving birth to a child. This obviously allows solving the attrition problem by considering it as a non-informative right censure problem for the duration (timing at first childbirth).

Since we consider a simple probit without random effects, the presence of unobserved heterogeneity across women can result in inconsistent estimation. However, it seems that the β parameters are not very sensitive to the omission of unobserved random effects as long as

[5] We have also tried different specifications for the binary model, in particular the complimentary log-log and the logit models. Results do not seem to be sensitive to changes in the distributional assumption.

the random effects are uncorrelated with the explanatory variables and flexible duration dependence is allowed. This result is explicitly suggested by Dolton and Van der Klaauw (1995) and confirmed by several other empirical findings, see for example Trussell and Richards (1985).

Notice that the distribution of the timing at first motherhood can be defective if there are women who decide to remain childless. Nevertheless, we assume that the probability that young women decide to remain childless for the rest of their life is not very high. Most of young women do not probably take such a drastic decision at the beginning of their reproductive career; they just postpone the decision to the future. It is obvious that the consequence of continuous postponement may be an increase in the risk to remain childless, because of possible fecundity impairments. Since it is not easy to distinguish between the two types of childlessness (a deliberate choice or as a consequence of continuous postponement), we assume that women do not refuse motherhood deliberately when they are young. In this way, we can assume a non-defective distribution for the timing at first birth.

We would like to emphasize that we are using the above described sequential probit model to conduct an exploratory analysis of the timing of first motherhood. We are aware of possible endogeneity problems due to interdependence between education, labour and fertility decisions and we do not pretend to explain causal relationships with our model. Nevertheless, our model for the transition to motherhood is consistently estimated if we consider sequential probit models for the the timing of leaving school and starting first job and if the endogeneity of education and work decisions are due to correlation between contemporaneous error terms in the different sequential models. The probit specification implies that the propensity to give birth at age t for a woman childless at age (t-1), say $r^*_{i,t}$, is a linear function of the explanatory variables plus and error term, i.e. $r^*_{i,t} = X_{i,t}\beta + \varepsilon_{i,t}$, and

$$\Pr(r_{i,t} = 1 \mid r_{i,t-1} = 1, X_{i,t}) = \Pr(r^*_{i,t} > 0) = \Pr(\varepsilon_{i,t} > -X_{i,t}\beta) = \Phi(X_{i,t}\beta) .$$

By assuming an analogous sequential probability model for the timing of completion of education, the propensity to complete education at age t for a woman who was studying at age (t-1), say $e^*_{i,t}$, is also linear function of the explanatory variables plus and error term, say $u_{i,t}$. If $\varepsilon_{i,t}$ and $u_{i,t}$ are distributed jointly as a bivariate normal with zero means, unit variances and covariance $\rho_{\varepsilon u}$; then the estimation of the sequential probit model for the first child timing neglecting the model for leaving school is still consistent. By analogy assuming the same type of sequential model for the propensity of transition to the first job at age t for a woman who never worked at age (t-1), and a similar structure for the error term, our model estimation is still consistent even neglecting both the models for leaving school and for beginning of the first job.

For a joint specification of hazard models considering instead individual random effects correlated between hazard models, we refer to Lillard (1993), Lillard et al (1995) and Upchurch et al (2002). We do not follow this joint specification approach, say Lillard's approach, for three main reasons. First, the Lillard's approach considers continuous rather than time-discrete models and requires observing multiple transitions to the same event for the same individuals which is not possible when we consider transitions to first time motherhood, first time worker and final completion of education. Second, the decomposition analysis of hazard rates at different ages would not be possible. Third, although the Lillard's approach controls for endogeneity by allowing dependence between individual random effects in different hazard models, it does not allow any correlation between error components which are time varying (see Upchurch et al 2002 for a further description of these approach's limits). In other words the Lillard's approach gives inconsistent results if the correlation between contemporaneous error terms is different from zero. Our model instead does impose a zero correlation between individual random effects but it still consistent if the contemporaneous error terms are correlated. Correlation between contemporaneous error terms occurs when there are temporary shocks affecting at the same time two decisions, in particular in our case the decisions to leave school, to become mother and to begin a first job. Any type of temporary shock in a woman behaviour affecting her probability of an involuntary pregnancy is likely to affect both the decisions of having a first child and leaving school. A temporary shock, like a temporary financial problem in the woman's household, can instead affect simultaneously the decisions to leave school, to start a first job and to give birth to a child.

3.2 Main findings

In this section we report the estimation results of the discrete-time hazard model for the 10 EU countries considered. Tables 3 to 5 show coefficients and p-values for each of the explanatory variables reported by column and each of the countries reported by row.

As expected, the hazard function is age dependent, it is first increasing and then decreasing in age virtually in all countries, but in Ireland where, however, the coefficients for the variables *age* and age^2 are not significantly different from 0, as shown in Table 3. The transition probability to first motherhood begins to decrease when women are in their thirties: this can be partly caused by fertility impairments rather than by a conscious decision to remain childless.

The age at the completion of the highest level of education (*age at hle*) has a negative effect on the probability to have a first child, but in Denmark where the education level does not have a significant effect. In other words, women with a higher level of education have a higher propensity to postpone their first childhood.

The hazard function is first increasing and then decreasing in the number of years since the completion of the highest level of education (*Age-Age at hle*) and in the number of years since the first work experience (*Age-Age at 1st job*), see Tables 4 and 5.

However, in Belgium, Germany, Greece and Ireland the number of years since the completion of the highest level of education is not very important once controlled for the age at the completion highest level of education and the number of years since the first work experience. Women wait to give birth to their first child on average from less than 2 (in Germany) to more than 7 years (in UK and in Italy) since the completion of education, and from less than 3 (in Greece) to about 7 years (in Denmark, Ireland and the UK) since the entry into the labour market. The effect of being still in education is negative except in Denmark where, however, the coefficient is not significantly different from *0* and where women begin usually to work before completing their education. The degree of the inhibiting effect of being a student differs across Europe.

In Figure 1 we report the non-parametrically estimated hazard and survival function for each country. The hazard function is smoothed by using a kernel function, while the survival function is simply based on the Kaplan and Meier (1958) nonparametric estimator. It seems evident that in Italy, Greece, Spain and Portugal the hazards functions are quite low at all ages, whereas the highest hazard function is observed in Denmark. We would like to investigate whether these differences are due to cross-country variations of age at first job and at the completion of the highest level of education. A first way to investigate that is by computing and comparing across countries the hazard function for different typologies of women. More precisely, using the estimated coefficients of the duration models we predict two hazard profiles for the following two typologies of women:

1. profile woman A, who is supposed to complete her highest level of education at 18 years old and to begin working at 20 years old,

2. profile woman B, who is supposed to complete her highest level of education at 23 years old and to begin working at 25 years old.

In Figures 2 we compare the hazard profiles for women A and B. Notice that the age at highest level of education and the age at the first job experience allow us to know all the explanatory variables used in our hazard model. In other words, considering the two typologies of women is equivalent to conditioning to specific values for the explanatory variables. If the differences in the hazard function are due to a genuine different propensity to motherhood between countries, then we should observe different profiles between countries when comparing women A, B. Looking at the profiles it seems evident that there are two groups of countries: one with a genuinely higher motherhood propensity, which is given by Belgium, Denmark, France and Germany, and one with a lower motherhood propensity, which is given instead by Greece, Ireland, Italy, Portugal, Spain and the UK.

Differences between countries in the hazard function profiles for women A and B are due only to differences in the estimated coefficients because we consider women perfectly identical in terms of work and education related variables. It seems therefore that the country with the highest propensity to motherhood, once controlled for the explanatory variables considered in our model, is France and not Denmark. The higher nonparametric hazard function observed for Denmark in Figure 1 seems therefore to be due to a difference in the distribution of the women characteristics more than to a general higher motherhood propensity.

Looking at Figure 2 we notice that, in general, delaying the completion of the highest level of education and the entry into the labour market implies a delay of the first childbirth as well as an increase in the probability to remain childless for all countries. In all countries, women B who complete their education at 23 and begin to work at 25 are less likely to give birth to a child and more likely to postpone motherhood then women A who complete their education at 18 and begin to work at 20. It seems moreover that in Greece, Italy and Portugal women who complete their education later, say at 23, have a decrease in the hazard function during the unemployment or inactive period before beginning the first job[6].

To investigate better whether the differences in the hazard functions between countries are due to variations in labour participation, timing in completing the education and timing in entering into the labour market, we conduct a decomposition analysis in the following section.

4. Explaining differences across countries

In this section we decompose the differences in the observed transition probabilities, $\Pr(r_{i,t}=1|r_{i,t-1}=0)$, between Italy and each of the other EU countries into two additional components:

(1) a component due to differences in the distribution of the women variables, say compositional component;

(2) a residual component due to differences in the impact of the variables on the propensity to give birth to a first child.

In other words we try to understand whether the difference between the hazard rate of two countries is explained by a genuine difference in the propensity to give birth to a first child for women living in two different countries or whether instead it is a consequence of different education and job experiences, which a woman has to face being living in different countries.

[6] It should be noted that we are interpreting these estimates as if they came from a real cohort. Actually, we describe the behaviour of a plurality of cohorts. In this sense it is possible that younger women who have never entered the labour market are simply not employed yet, while the older ones are the ones who made a clear family-oriented choice.

By hazard rates (or probability of transition) we mean the probability to have a first child for childless women at a specific age. In particular we consider the hazard rates for women 19, 22, 28 and 34 years old. Notice that the transition probabilities are computed considering our sample, which is not representative of a specific cohort, year or age.

Demographers and sociologists have been the first to decompose differences in rates between two populations. We refer to Kitagawa (1955) for a first thorough explanation on how to decompose rate differences in two components and authors cited therein for earlier references. Kitagawa (1955) clarifies for the first time the link between standardized rates and decomposition of differences between two rates.

Demographers and sociologists estimate usually nonparametrically conditional rates by using the empirical relative frequencies for all possible realizations of the explanatory variables. Unfortunately, nonparametric estimation may perform very poorly when the number of the explanatory variables is high. More recently economists have instead begun to decompose total rates by considering parametric probability models, which allow for a larger set of variables, both continuous and categorical, characterizing the populations to be compared[7]. The parametric approach involves the estimation of a parametric probability model for each population to be compared. In our case we use the probit model already introduced to estimate the transition probabilities to motherhood for each of the 10 EU countries considered.

4.1 Description of the decomposition method

Let us consider again the simple probit model for the timing at the first childbirth

$$\Pr(r_{it} = 1 | r_{it-1} = 0, X_{i,t}, t) = \Phi(\beta X_{i,t} + a + b\ t + c\ t^2),$$

[7] See Goumulka and Stern (1990) for a first example of the decomposition analysis applied to unemployment probit model.

and let the coefficient vectors β and γ=[a,b,c] vary across countries. Let β_1 and $\gamma_1 = [a_1, b_1, c_1]$ be the coefficients for Italy and β_0, $\gamma_0 = [a_0, b_0, c_0]$ the ones for a specific EU country with which we are interested to make a comparison. Let $f_o(X_{i,t})$ and $f_1(X_{i,t})$ be the density distribution functions for the country *0* (EU comparison country) and *1* (Italy), and d_i^0 and d_i^1 be two dummy variables indicating if a generic woman *i* belongs to country *0* or to country *1*. Integrating out the explanatory variables, *X*, from the hazard rate we obtain the transition probability from $r_{i,t-1}$=0 to $r_{i,t}=1$. The transition probability for country *0* is given by:

$$\Pr(r_{i,t} = 1 | r_{i,t-1} = 0, d_i^0 = 1) = \int \Pr(r_{i,t} = 1 | r_{i,t-1} = 0, d_i^0 = 1, X_{i,t}, t; \beta_0, \gamma_0) f_0(X_{i,t}) dX_{i,t} \ ,$$

and the analogous probability for the country *1* is given by

$$\Pr(r_{i,t} = 1 | r_{i,t-1} = 0, d_i^1 = 1) = \int \Pr(r_{i,t} = 1 | r_{i,t-1} = 0, d_i^1 = 1, X_{i,t}, t; \beta_1, \gamma_1) f_1(X_{i,t}) dX_{i,t} \ .$$

The difference between the two above transition probabilities can be decomposed in the following way:

$$\Pr(r_{i,t} = 1 | r_{i,t-1} = 0, d_i^1 = 1) - \Pr(r_{i,t} = 1 | r_{i,t-1} = 0, d_i^0 = 1) =$$

$$= \int \Pr\left(r_{i,t} = 1 | r_{i,t} = 0, X_{i,t}, t; \beta_1, \gamma_1\right) \left[f_1(X_{i,t}) - f_0(X_{i,t})\right] dX_{i,t} +$$
$$+ \int f_0(X_{i,t}) \left[\Pr\left(r_{i,t} = 1 | r_{i,t-1} = 0, X_{i,t}; \beta_1, \gamma_1\right) - \Pr\left(r_{i,t} = 1 | r_{i,t-1} = 0, X_{i,t}; \beta_0, \gamma_0\right)\right] dX_{i,t}$$

The last equation shows how the difference between the transition probabilities, observed for two different countries, can be decomposed into two components. The first component, given by the first addend in the right hand side of the last equation, represents differences in the transition rates due to a different composition of the populations. By differences in the population composition we mean differences in the distributions of the explanatory variables. The second component - the residual component - given by the second addend in the right hand side, represents the "genuine" difference in the transition rates after controlling for the specific set of explanatory variables used. Since we use a parametric model for the transition probabilities, the residual component may be also defined as the effect of changes in the parameters, differences between (β_0, γ_0) and (β_1, γ_1).

The transition probability for a specific country, say *0* (or *1*), can be estimated just by replacing the coefficients β_0 and γ_0 (β_1 and γ_1) with their estimates and by considering the sampling average instead of the integral in the following way

$$\hat{p}_o = \sum_i \frac{d_i^0 \Pr(r_{i,t} = 1 \big| r_{i,t-1} = 1, X_{i,t,} t; \hat{\beta}_0, \hat{\gamma}_0)}{\sum_i d_i^0},$$

$$\hat{p}_1 = \sum_i \frac{d_i^1 \Pr(r_{i,t} = 1 \big| r_{i,t-1} = 1, X_{i,t,} t; \hat{\beta}_1, \hat{\gamma}_1)}{\sum_i d_i^1},$$

where the $\sum_i$ is over all individuals belonging to the countries *0* and *1*.

The two terms of the decomposition can be instead estimated as follow:

$$\hat{p}_1 - \hat{p}_o = \left(\sum_i \frac{d_i^1 \Pr(r_{i,t} = 1 \big| r_{i,t-1} = 1, X_{i,t,} t; \hat{\beta}_1, \hat{\gamma}_1)}{\sum_i d_i^1} - \sum_i \frac{d_i^0 \Pr(r_{i,t} = 1 \big| r_{i,t-1} = 1, X_{i,t,} t; \hat{\beta}_1, \hat{\gamma}_1)}{\sum_i d_i^0} \right)$$
$$+ \left(\sum_i \frac{d_i^0 \Pr(r_{i,t} = 1 \big| r_{i,t-1} = 1, X_{i,t,} t; \hat{\beta}_1, \hat{\gamma}_1)}{\sum_i d_i^0} - \sum_i \frac{d_i^0 \Pr(r_{i,t} = 1 \big| r_{i,t-1} = 1, X_{i,t,} t; \hat{\beta}_0, \hat{\gamma}_0)}{\sum_i d_i^0} \right).$$

Since we would like to evaluate the effect on the Italian hazard rates of a change of the distribution of the women's characteristics keeping the Italian intrinsic childbirth propensities constant, we need to define the compositional component as the effect of a change of the Italian population composition with the one of another EU country but keeping the hazard rates coefficients estimated for Italy. This implies a residual component computed using the distribution of the characteristics of the country 0, which is the specific EU country compared with Italy.

4.2 Results of the decomposition

We present the results of the estimation of the transition rate decomposition in Tables 6-9. We look at differences between Italy and all other countries in our sample (Belgium, Denmark, France, Germany, Greece, Ireland, Portugal, Spain and the UK).

The choice of Italy as a point of reference for our comparisons is determined by three main reasons. First of all we point out a statistical motive: the Italian sample is the largest. A second reason is that Italy represents a puzzling case of low fertility, late birth regime and low female labour participation compared to the other countries (Bettio and Villa 1998). The last reason – strictly linked to the second – concerns policy implications. In fact, we are interested in evaluating the possible effects of a different structural context on entry into motherhood. For instance, we want to evaluate whether different degree of labour participation and a more precocious timing of leaving school and of entry into the labour market might affect an earlier childbirth.

We present the results of the decomposition of the differences in the transition probability from childlessness to motherhood at age 18, 22, 28 and 34 respectively in Tables 6-9. The results of the decompositions at different ages change because the distribution of the women characteristics varies at different ages.

The results for first birth rates at 22 and 28 are very similar. Differences between Italy and Ireland, Portugal, Spain and the UK are mainly due to difference in the women characteristics whereas differences between Italy and all other countries are mainly due to coefficients. It should be noted, however, that the differences in the hazard rate due to population structures – even if not dominant - are even more relevant (in absolute term) for France and Germany at the age of 22, and for Belgium, Denmark and France at the age of 28.

Different results are instead observed for the probability of transition to motherhood for childless women at 18 and at 34. For childless women at 18 we can observe a much lower first birth rate than women at 22, 28 and 34 for all countries, as expected. Moreover, the differences between countries in the first birth rates at 18 are very low and never higher than 2 percentage points. Explaining those differences is therefore not very relevant.

Childless women at 34 years have very different probabilities of transition to motherhood across countries, from 2.8% for Portugal to 9.5% for Denmark. Nevertheless, in six out of nine countries the difference is mainly due to relevant differences in the characteristics rather than genuine differences in the women propensity. Older childless women seem to have a more similar propensity to motherhood especially once controlled for their characteristics. The only exceptions are France, Greece and Portugal. Everything being equal, childless women at 34 years old seem to have a slight lower propensity to childhood – even lower than in Italy - in Greece and Portugal, and a higher propensity instead in France.

When the transition rate differences are mainly explained by differences in the coefficients, policy interventions to modify the labour market or the education system would have a doubtful effect in increasing the low Italian transition rate up to the same level of other EU countries. The main part of the difference, in fact, would remain. Nevertheless, policies may have some effect in increasing the hazard rate towards the one observed in other countries, where the differences due to variables are conspicuous, even if not prevalent: this might be the case of Belgium, Denmark and France at the age of 28, for instance.

In summary, it seems that, if we were able to change the work and education characteristics of Italian women toward the ones observed for Ireland, Portugal, Spain and the UK (where the compositional effect is dominant), it would be possible to narrow the differences in the first birth rates. Admittedly, a movement of the Italian first birth rate toward the Portuguese and Spanish ones would not imply a big increase. It is definitely more interesting instead considering a movement of the Italian rate closer to the higher rates observed for Ireland and the UK.

The lesson from these results is that Italy can learn something on how to increase its first birth rates by looking at women's characteristics in other countries, especially Ireland and the UK. However, little can be learnt from comparison with countries which are far away from Italy, such as Germany, Denmark, France and Belgium, where the propensity to motherhood for childless women is much higher even after controlling for differences in the women characteristics.

Looking at the descriptive statistics (Table 1), it seems that Italian women enter into the labour market later than the British and Irish women, moreover the waiting time between leaving school and beginning the first job is on average of 1.5 years for Italy and less than 2 months for Ireland and the UK. Policies oriented to reduce the unemployment duration of young women looking for first job could then have an effect in speeding up both the entry into the labour market and the possible transition to motherhood.

5. Conclusions

Using data from the 8 waves of the ECHP, this paper estimates a duration model for the timing at first birth for 10 European countries, namely Belgium, Denmark, France, Germany, Greece, Ireland, Italy, Portugal, Spain and the UK. In most countries, we find that higher levels of education have in general a double effect on the first birth event: a postponement of it and a reduction of its probability. In all countries we find a very strong relationship between the timing at first birth and the age at the beginning of the work career. Women, after the beginning of their first job, wait on average between 3 and 7 years before deciding to have their first child. It seems that there are fewer housewives consecrating their life to childcare with respect to the past, whereas there are more women with higher level of education, career-oriented and not keen on having a first child at the very beginning of their career. Our results provide also empirical evidence for the existence of a biological age constraint for fertility. As expected, the probability to have a first child tends to increase with age until about 30 years old and then tends to decrease.

There are probably mainly two types of policies which can have an impact on the first child timing: family friendly policies and tempo policies. By family-friendly policies we mean policies aiming at reconciling motherhood and work by increasing for example formal childcare facilities and giving incentives for flexible working arrangements (e.g. part-time). By tempo policies we mean instead policies specifically designed to solve the motherhood postponement problem, by for example shortening the average school duration and reducing young women's unemployment to speed up their entry into labour market.

By decomposing differences between first birth hazard rates in two components, it is possible to some extent to evaluate the impact of these policies. The compositional component is linked to differences in the distributions of the timing of leaving school and starting first job; therefore differences between countries in tempo policies should reflect in differences in this component. Whereas the residual component is linked to any residual unobserved heterogeneity between countries, in particular this may be associated to differences in cultural context, in family-friendly-policies but also in anything else we have not control for in the estimation of our hazard models. Therefore, even if it is plausible to assume that big residual components, observed when comparing for example Denmark and Italy, are - at least in part - due to differences in family-friendly policies, we do not push further this interpretation of the results in this paper. We focus instead our attention of the effect of tempo policies whose potential effect is more undeniably linked to the compositional component due to differences in the characteristics.

From our analysis we find that if Italian women were experiencing the same work and education patterns than in Ireland, Portugal, Spain and the UK, then the low first birth hazard rate at age 22 and 28 for Italy would get closer to the higher rates observed for those countries. This means that policies aiming, for example, at reducing the delay in the first job experience for Italian women with respect to the above four countries can reduce the gap in the first birth hazard rates at 22 and 28 between Italy and those counties. The efficacy of such policies is instead questionable when trying to reduce the gap between Italy and the remaining 5 countries, where the hazard rates are different even after controlling for the women characteristics. Yet, there are countries where the tempo policies could have an important impact because the compositional effect although not dominant is conspicuous (for example Belgium, Denmark and France for the hazard rates at age 28).

Considering the first birth hazard rates for older women aged 34, the gap between Italy and other EU countries is mainly due to differences in the characteristics in 6 out of 9 cases. Only Greece, Portugal and France seem to differ from Italy because of a genuine different motherhood propensity or because of unobserved characteristics we were not able to control for.

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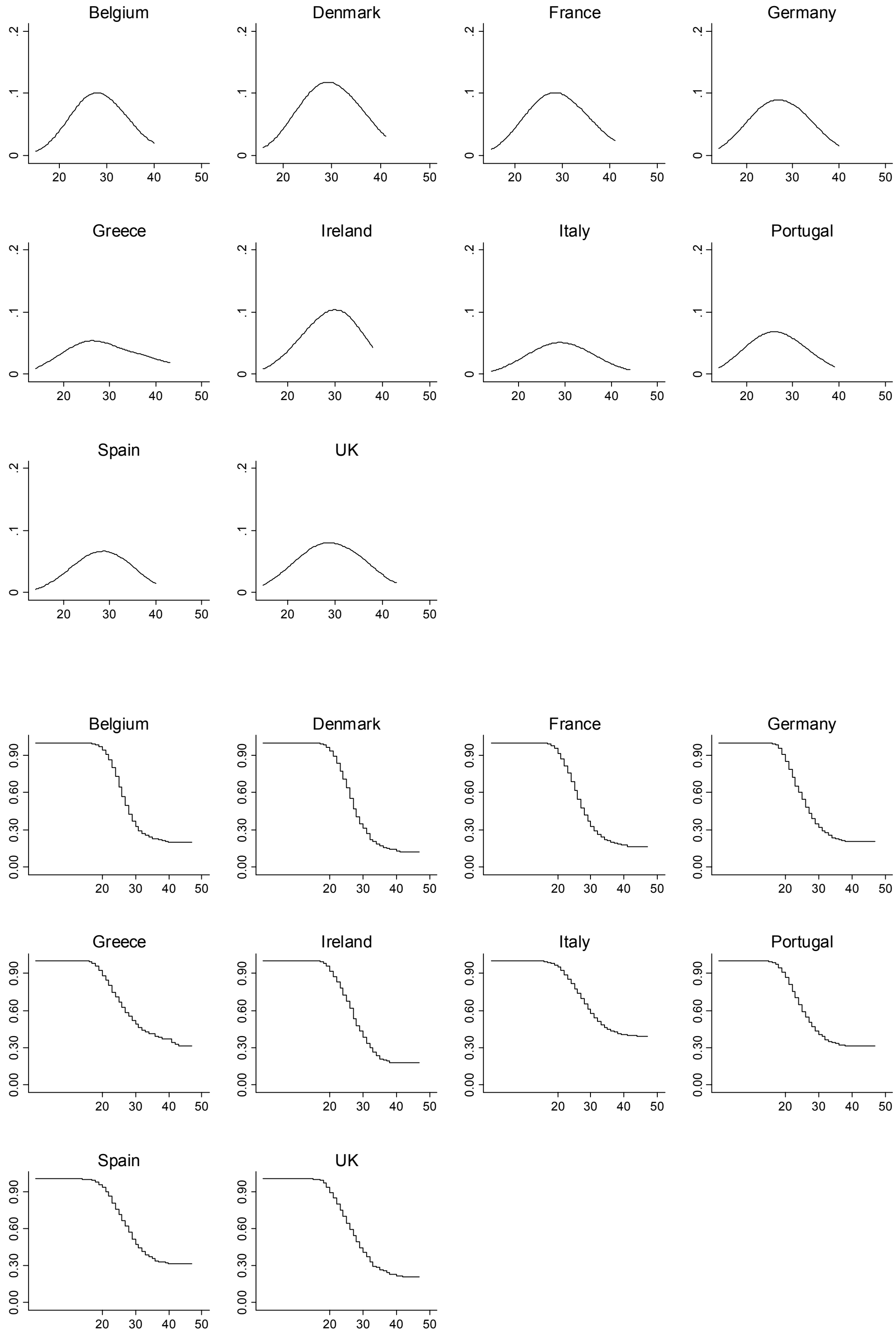
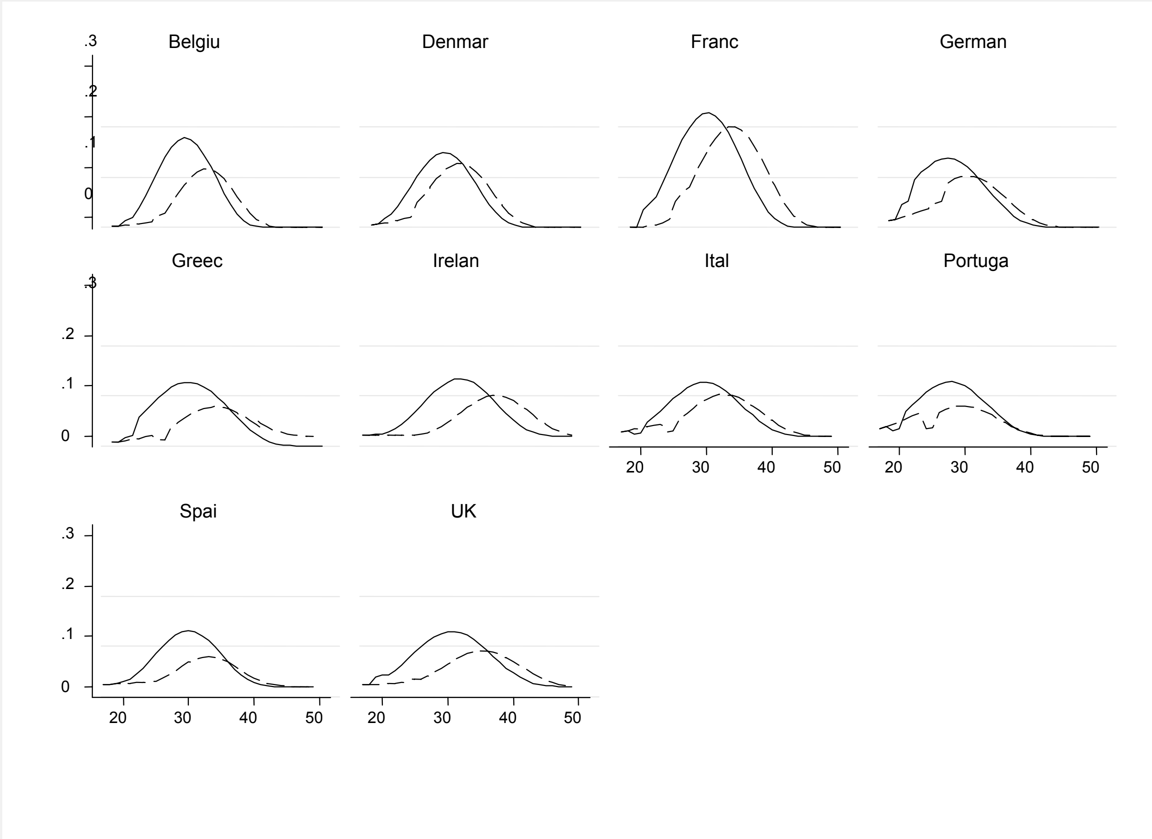
Tables and figures

Figure 1. First childbirth hazard and survival function estimated nonparametrically[8]



[8] We report the estimated hazard and the survival functions estimated by country. The hazard function is smoothed by using a kernel function. The Kaplan-Meier estimator is used instead for the survival.

Figure 2. First childbirth hazard rates for woman A (solid line) and B (dashed line)[9]



[9] Woman A is supposed to complete her highest level of education at 18 years old and to begin working at 20 years old, whereas woman B is supposed to complete her highest level of education at 23 years old and to begin working at 25 years old.

Table1 Sample size, percentage of women still childless when last observed, and average and standard deviation of their age

Country	N. of women	N. women still childless when last observed	Percentage women childless when last observed	Average (S.D.) age for women childless when last observed
Belgium	1395	479	34.34%	29.94 (6.94)
Denmark	1282	424	33.07%	27.33 (6.86)
France	2089	530	25.37%	27.45 (5.73)
Germany	1261	491	38.94%	27.65 (7.27)
Greece	2717	1516	55.80%	27.32 (6.64)
Ireland	2089	1010	48.35%	25.83 (5.55)
Italy	4126	2572	62.34%	27.54 (6.85)
Portugal	2220	1032	46.49%	27.84 (7.32)
Spain	3999	2212	55.31%	26.97 (6.69)
UK	2043	781	38.23%	29.06 (7.63)

Table 2 Summary statistics for age at 1st child, 1st job and at highest level of education.

Country	Age at 1st child (S.D.)	Age at 1st job (S.D.)	Never worked	Age at hle (S.D.)
Belgium	25.61 (3.88)	20.64 (3.29)	0.08	20.32 (3.74)
Denmark	25.63 (4.17)	18.80 (2.86)	0.05	23.06 (4.06)
France	24.47 (3.92)	19.06 (2.76)	0.11	18.22 (3.08)
Germany	23.60 (4.29)	20.19 (3.81)	0.03	22.25 (5.41)
Greece	23.85 (4.56)	21.23 (4.55)	0.30	17.69 (4.02)
Ireland	24.89 (4.38)	18.10 (2.33)	0.08	18.60 (3.59)
Italy	25.59 (4.62)	20.75 (4.80)	0.34	18.45 (5.32)
Portugal	23.54 (4.28)	18.90 (5.02)	0.21	16.73 (5.49)
Spain	25.03 (4.36)	19.20 (4.66)	0.21	18.80 (5.33)
UK	25.28 (4.89)	18.24 (3.42)	0.08	17.93 (3.14)

Note: The averages and standard deviations (reported in parenthesis) are computed considering one single observation for each woman in our sample. Age at 1st child, 1st job and at hle (highest level of education) are retrospective information, while the dummy for women who never worked refers to the wave when women are last observed in our sample. The averages of age at 1st child, 1st job and at highest level of education are computed excluding women still childless, without any work experience and still at school.

Table 3. Hazard model estimated by country: intercept and age coefficients.

Country	Constant	p-value	Age	p-value	Age^2	p-value	N.	-Log-L	LR test[a]	p-value
Belgium	-5.67	0.000**	0.345	0.000**	-0.006	0.000**	16154	-2981.6	1005.8	0.000
Denmark	-7.23	0.000**	0.384	0.000**	-0.007	0.000**	10856	-2114.8	596.8	0.000
France	-6.417	0.000**	0.355	0.000**	-0.006	0.000**	30252	-5647.1	1959.9	0.000
Germany	-4.21	0.000**	0.298	0.000**	-0.006	0.000**	29183	-6314.1	827.3	0.000
Greece	-1.821	0.000**	0.116	0.001**	-0.001	-0.056	26884	-4204.2	1003.6	0.000
Ireland	-1.198	0.042*	-0.065	-0.217	0.001	-0.470	20320	-3525.3	1075.9	0.000
Italy	-4.683	0.000**	0.201	0.000**	-0.004	0.000**	45415	-5843.3	1433.5	0.000
Portugal	-4.981	0.000**	0.3	0.000**	-0.006	0.000**	22229	-4025.0	831.7	0.000
Spain	-5.09	0.000**	0.241	0.000**	-0.005	0.000**	40409	-6097.0	1585.7	0.000
UK	-2.751	0.000**	0.102	0.003**	-0.002	0.004**	23703	-4678.1	725.2	0.000

Note: The table reports coefficients and p-values for the variables reported by column and the countries reported by row. 2 asterisks and 1 asterisk indicate significance of the coefficients at 1% and 5% level. The number of observation, N, used for each country specific model, the likelihood ratio test and its p-value are also reported.

[a] LR test is the likelihood ratio test for the joint significance of the full set of explanatory variables: age, age^2, age at hle (age when the highest level of education was completed), Age-Age at hle, $(\text{Age-Age at hle})^2$, Still in education dummy, Age-Age at 1st job, $(\text{Age-Age at 1st job})^2$, dummy for women who never worked.

Table 4. Hazard model estimated by country: education variables coefficients.

Country	Age at hle[a]	p-value	(Age-Age at hle)	p-value	$(\text{Age-Age at hle})^2$	p-value	Still in education	p-value
Belgium	-0.052	0.000**	0.039	-0.133	-0.005	0.000**	-1.259	0.000**
Denmark	0.017	0.313	0.093	0.000**	-0.005	0.000**	0.129	0.752
France	-0.02	0.014*	0.087	0.000**	-0.006	0.000**	-0.882	0.000**
Germany	-0.052	0.000**	0.011	-0.601	-0.003	0.057	-1.283	0.000**
Greece	-0.095	0.000**	0.015	-0.503	-0.005	0.000**	-2.216	0.000**
Ireland	-0.026	0.088	0.047	-0.053	-0.003	0.007**	-0.684	0.058
Italy	-0.008	0.372	0.073	0.000**	-0.004	0.000**	-0.036	0.863
Portugal	-0.033	0.001**	0.052	0.003**	-0.003	0.000**	-0.567	0.025*
Spain	-0.023	0.008**	0.063	0.000**	-0.003	0.000**	-0.421	0.047*
UK	-0.044	0.000**	0.045	0.010**	-0.003	0.000**	-0.97	0.000**

Note: The table reports coefficients and p-values for the variables reported by column and the countries reported by row. 2 asterisks and 1 asterisk indicate significance of the coefficients at 1% and 5% level.

[a] Age at hle means age at which the highest level of education was completed.

Table 5. Hazard model estimated by country: work relates variables coefficients.

Country	Age-Age at 1st job	p-value	(Age-Age at 1st job)^2	p-value	Never worked [a]	p-value
Belgium	0.138	0.000**	-0.005	0.000**	-0.108	0.253
Denmark	0.049	0.032*	-0.001	0.267	0.009	0.937
France	0.093	0.000**	-0.004	0.000**	0.113	0.065
Germany	0.057	0.000**	-0.003	0.000**	-0.176	0.000**
Greece	0.071	0.000**	-0.002	0.050*	-0.34	0.000**
Ireland	0.236	0.000**	-0.008	0.000**	-0.025	0.843
Italy	0.062	0.000**	-0.002	0.011*	-0.361	0.000**
Portugal	0.037	0.005**	0.000	0.690	-0.38	0.000**
Spain	0.126	0.000**	-0.004	0.000**	0.04	0.507
UK	0.115	0.000**	-0.004	0.000**	0.141	0.071

Note: The table reports coefficients and p-values for the variables reported by column and the countries reported by row. 2 asterisks and 1 asterisk indicate significance of the coefficients at 1% and 5% level.
[a] Never work is a dummy taking value 1 is a woman never worked, but it takes value 0 if a woman still have to complete her education.

Table 6. Differences in the first birth hazard rates for childless women at 18

Country	Difference in hazard rates			Hazard rate
	Total	Due to variables	Due to coefficients	
Belgium	0.001	0.006	-0.005	0.009
Denmark	0.002	0.006	-0.004	0.008
France	-0.003	0.007	-0.010	0.013
Germany	-0.018	0.020	-0.038	0.028
Greece	-0.016	0.001	-0.017	0.026
Ireland	-0.005	-0.002	-0.003	0.015
Portugal	-0.019	-0.010	-0.008	0.029
Spain	-0.003	-0.003	0.000	0.013
UK	-0.013	-0.001	-0.013	0.023

Note: The table reports the differences in first birth rates between Italy and each of the country indicated by row. Moreover in last column the hazard rates, i.e. the first birth rates for childless women at 18, are reported.

Table 7. Differences in the first birth hazard rates for childless women at 22

Country	Difference in hazard rates			Hazard rate
	Total	Due to variables	Due to coefficients	
Belgium	-0.021	0.006	-0.027	0.053
Denmark	-0.014	0.011	-0.025	0.046
France	-0.032	0.019	-0.050	0.064
Germany	-0.044	0.022	-0.066	0.076
Greece	-0.016	0.005	-0.022	0.049
Ireland	-0.019	-0.021	0.002	0.052
Portugal	-0.033	-0.015	-0.018	0.065
Spain	-0.011	-0.010	-0.001	0.043
UK	-0.024	-0.013	-0.011	0.056

Note: The table reports the differences in first birth rates between Italy and each of the country indicated by row. Moreover in last column the hazard rates, i.e. the first birth rates for childless women at 22, are reported.

Table 8. Differences in the first birth hazard rates for childless women at 28

Country	Difference in hazard rates			Hazard rate
	Total	Due to variables	Due to coefficients	
Belgium	-0.069	-0.038	-0.031	0.129
Denmark	-0.075	-0.033	-0.042	0.135
France	-0.058	0.032	-0.091	0.118
Germany	-0.044	-0.006	-0.037	0.103
Greece	-0.002	0.003	-0.005	0.061
Ireland	-0.048	-0.057	0.010	0.107
Portugal	-0.015	-0.009	-0.006	0.074
Spain	-0.017	-0.016	-0.001	0.076
UK	-0.034	-0.033	-0.001	0.093

Note: The table reports the differences in first birth rates between Italy and each of the country indicated by row. Moreover in last column the hazard rates, i.e. the first birth rates for childless women at 28, are reported.

Table 9. Differences in the first birth hazard rates for childless women at 34

Country	Difference in hazard rates			Hazard rate
	Total	Due to variables	Due to coefficients	
Belgium	-0.025	-0.032	0.007	0.061
Denmark	-0.058	-0.055	-0.004	0.095
France	-0.025	0.007	-0.032	0.062
Germany	-0.006	-0.018	0.012	0.043
Greece	0.002	0.000	0.002	0.034
Ireland	-0.046	-0.047	0.001	0.082
Portugal	0.009	-0.004	0.013	0.028
Spain	-0.008	-0.014	0.006	0.045
UK	-0.030	-0.030	-0.001	0.067

Note: The table reports the differences in first birth rates between Italy and each of the country indicated by row. Moreover in last column the hazard rates, i.e. the first birth rates for childless women at 34, are reported.