DIFFERENCES IN DELAYING MOTHERHOOD ACROSS EUROPEAN COUNTRIES: EMPIRICAL EVIDENCE FROM THE ECHP

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ABSTRACT

Age at motherhood has increased in most European Countries in the past decades. The main aim of this paper is to assess the impact of women's education and work experience on the timing of first birth across the European Union (EU).

According to the literature - based on income maximisation framework (Gustafsson 2001, Hotz et al. 1997) - women with a higher degree of education 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, higher educated women or career women seem to enter motherhood earlier in the Northern European Countries (Kravdal 1994, Hoem 2000, Andersson 2001). Conceivably, these ambiguous findings might reflect substantial cross-country differences that we would like to point out. Therefore, we conduct an analysis to explain how the probability to enter into motherhood differs across 10 European Union countries by using the European Community Household Panel survey (ECHP).

On one side, the gap between countries may reflect differences in the observed characteristics of the national women populations, such as differences in the female labour participation and in the human capital investment. On the other side, the gap may be instead due to different fertility propensities across countries. In the empirical application we try to disentangle between these two reasons.

NON-TECHNICAL SUMMARY

One of the most characteristic features of fertility changes in Europe in last decades has been a general and progressive postponement of motherhood. Unfortunately, a first child delay is not completely costless. Late mothers are more exposed to risks for their own and their child's health. Moreover, due to biological age limits, women are more likely to need medical assistance, e.g. in vitro fertilization, and to risk an involuntary permanent childless.

In this paper we try to identify potential determinants of motherhood postponement in 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 age at first birth and the age at the beginning of the work career. It seems that there are fewer women 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.

Governments have been and will be pushed to take a clear position concerning fertility treatments to lengthen the women biological age limit. But a more ethic solution could be instead adopting policies to help women to speed up the first steps to adulthood and in particular the beginning of their work career. Our empirical analysis gives evidence that an early completion of education and an early entry into the labour market are associated with an early entry into motherhood in all European countries. The question is then whether policies aimed at speeding up the first steps to adulthood can halt, or at least reduce, motherhood postponement.

To answer to this question we compare women between countries. In particular we focus attention on Italy, where the postponement of the first childbirth is particularly evident, and we compare it with other EU countries, namely Belgium, Denmark, France, Germany, Greece, Ireland, Portugal, Spain and the UK. We compute for each country the first birth rates for childless women at age 18, 22, 28 and 34. Differences in those first birth rates are due to two different causes: differences in the education and work characteristics between countries and differences in the intrinsic motherhood propensity between countries.

If the differences between Italy and the other countries were mainly due to differences in women's characteristics, policies favouring earlier completion of education and beginning of work in Italy would help to reduce the gap with the rest of Europe not only in this respect, but also with regard to the start of fertility. If, instead, differences were mainly due to intrinsic lower motherhood propensity in Italy, then policies would probably be scarcely effective.

From our analysis we find that first birth rates for women at 18 are very low, as expected, in all the 10 EU countries studied. Moreover, differences between countries are always lower than 2%. More interestingly, first birth rates for women at 22, 28 and 34 are instead higher and quite different across countries. Italian women at 22 and 28 have a lower first birth rate then in all other countries. The difference between Italy and each of the other countries is mainly due to different women characteristics for Ireland, Portugal, Spain and the UK, while it is mainly due to an intrinsically higher motherhood propensity in Belgium, Denmark, France, Germany, Greece. This mean 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 rates at age 22 and 28 for Italy would get closer to the higher rates observed for those countries.

Moreover, we find that childless women at age 34 have more similar intrinsic motherhood propensities, once controlled for women characteristics, across EU countries. Considering an Italian woman without children, 34 years old and with specific work and education characteristics and a woman with the same characteristics in Belgium, Denmark, Germany, Ireland, Spain or UK, we observe similar probabilities of a first childbirth. Therefore, we can say that being able to change the work and education patterns of Italian women toward their European peers would help in reducing the first childbirth rate gap at age 34.

1. INTRODUCTION

In the last decades a general and progressive delay of the first child birth 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, Billari and Ortega 2002).

What are the consequences of this continuous delay for fertility levels? On the one hand, demographers have shown that it causes a downward distortion on fertility indicators observed in the period (*tempo effect*), overemphasising the decline of total fertility rates (Bongaarts and Feeney 1998). On the other hand, the delay of the first child birth is considered a relevant cause of low 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).

At macro level, the evidence of the relation between postponement and low fertility is quite controversial (Frejka and Calot 2001). In some countries – as, for instance, Northern European countries and Belgium - the delay of entry into motherhood has a "pure" postponement effect without any consequence for fertility levels, as the *catching-up* process can be strong enough to balance the previous delay completely. In other countries – e.g. France and Switzerland - cohort fertility levels seem to be affected only slightly. Finally, in other cases – e.g. the Southern Europe – a consistent contraction of cohort fertility follows the postponement of first birth for the cohorts born after 1960.

Micro level analysis seems to confirm more clearly the evidence that shifting motherhood to later ages is associated with a reduction of completed fertility (Morgan and Rindfuss, 1999; Billari and Kohler 2000). Kohler, Skyitte and Christiansen (2001), using longitudinal data from a sample of Danish twins, find that for every year by which the first birth is deferred, there is a reduction of 3% in the number of children at the end of reproductive life. Kohler, Billari and Ortega (2002) estimate that for Italy and Spain the postponement effect is even higher: in fact, each year of delay implies a reduction of completed fertility by between 2.9 and 5.1%. A weaker effect, declining across cohorts, has been observed in Northern Europe and in the US (Morgan and Rindfuss 1999).

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, age at first employment. 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 2001). Conceivably, these ambiguous findings might reflect substantial cross-country differences, which we would like to point out.

The aim of this paper is twofold. Firstly, we assess the influence of women's human capital and work experiences on the timing of first birth in each country, using a discrete hazard model. Secondly, we focus our attention on Italy, where the postponement of the first childbirth is particularly evident, and we check whether this delay is due to a lower intrinsic childbearing propensity, everything being equal, or to different education and work characteristics of the Italian women. In other words, 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 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?

The answer 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. Differences in the coefficients reflect an intrinsic difference in the childbearing propensity, which persists even if everything else - at least everything observable - is equal.

We then decompose the differences between the hazards observed for Italy and each of the other European countries in the part due to coefficients and the part due to characteristics. If the difference in the hazards is mainly due to differences in the variables, policies can plausibly have an impact on the probability of being mothers. If the gap between hazards is instead mainly due to the coefficients, then the effect is more ambiguous. In other words, a difference mainly due to coefficients means that the two countries compared have a different childbearing propensity even if every variable is equal. This may be due to different cultural factors or to heterogeneity in other factors, which we are not able to control for.

The work is divided as follows. Section 1 is dedicated to review the theoretical approaches to motherhood postponement. In Section 2, we present the data and the variables used in the study. In Section 3, we describe the statistical model used to analyse the determinants of the timing at first birth and we present the results of the model estimation. Section 4 is devoted to explain the differences in the timing at first birth between Italy and any other country using a decomposition analysis. 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, Billari and Ortega 2002, Simò, Golsch and Steinhage 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 investment in human capital and 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 26,733 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] 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.

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,
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 determined by the results of a previous explorative analysis, revealing the absence of any impact of current activity status and income on the hazard of having a first birth. We would suggest two possible reasons why the current employment situation observed is not relevant in explaining the timing at the first childbirth: first, fertility decisions may be linked essentially to long term labour choices; second, having at least a first child is a wish shared by the majority of women, regardless their current occupation.

We do not consider the women's marital status because retrospective data on marital status are not available in the ECHP. If marital and child birth timing decisions can be explained by a joint sequential probit model (where error terms in the two decision models are correlated), then it is still possible to consistently estimate the marginal probit for the timing at first childbirth. In other words the univariate sequential probit presented in next section is not affected by endogeneity and/or omission of the marital status variable.

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.

[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.

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 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, Kenjoh and Wetzels 2001).

In Table 1 simple descriptive statistics are reported. It is immediately evident that Italy represents a peculiar case: 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%.

4. 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.

4.1 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, 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.

[3] For more details on discrete-time duration models we refer to Allison (1982), Yamaguchi (1991), 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.

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 cubic 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:

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

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 the random effects are uncorrelated with the explanatory variables and flexible duration dependence is allowed. This result is confirmed by several empirical findings, see Trussell and Richards (1985) and Dolton and Van der Klaauw (1995). 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.[6]

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[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.

[6] Our model for the transition to motherhood is consistently estimated is we assume sequential probit models for the transition probability from out to in labour force, from in to out of labour force and for leaving school, and

4.2 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.

---

if the endogeneity of education and work decisions are due to correlation between contemporaneous error terms in the transition models.

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[7].

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.

5. 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,

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

[7] 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.

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.

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 hazard rates for women 19, 22, 28 and 34 years old.. Notice that transition probabilities are computed considering our sample, which is not representative of a specific cohort, year or age. We consider also the decomposition of

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 and authors cited therein for earlier references.

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[8]. 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.

5.1 Description of the decomposition method

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

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

and let the coefficients 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

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

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

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

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

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

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:

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

5.2 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. 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. Childless women at higher ages 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 towards other EU countries. 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, 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 (Tab. 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.

6. 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.

In general, family-friendly policies aim at reconciling motherhood and work by increasing for example formal childcare facilities and giving incentives for flexible working arrangements (e.g. part-time). No special policies have been specifically designed instead to solve the motherhood postponement problem. Governments have been and will be pushed to take a clear position concerning fertility treatments to lengthen the women biological age limit. But a more ethic solution could be instead adopting policies to speed up the transition to adulthood for young women. Our results give evidence that an early completion of education and an early entry into the labour market are associated with an early entry into motherhood. The question is then whether policies aimed at speeding up the first steps to adulthood can halt, or at least reduce, motherhood postponement. The answer is not straightforward because we cannot be sure that women pushed to the adulthood beforehand be ready to give birth to a child, so that their propensity to motherhood could stay low and the delay syndrome could persist.

The answer to the above question would require a human experiment, which obviously is not possible. Comparison of countries with populations experiencing different work and education patterns can approximate such a natural experiment. The decomposition analysis of the first birth rates at different ages between Italy and other EU countries is then the statistical tool we used to find an answer.

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.

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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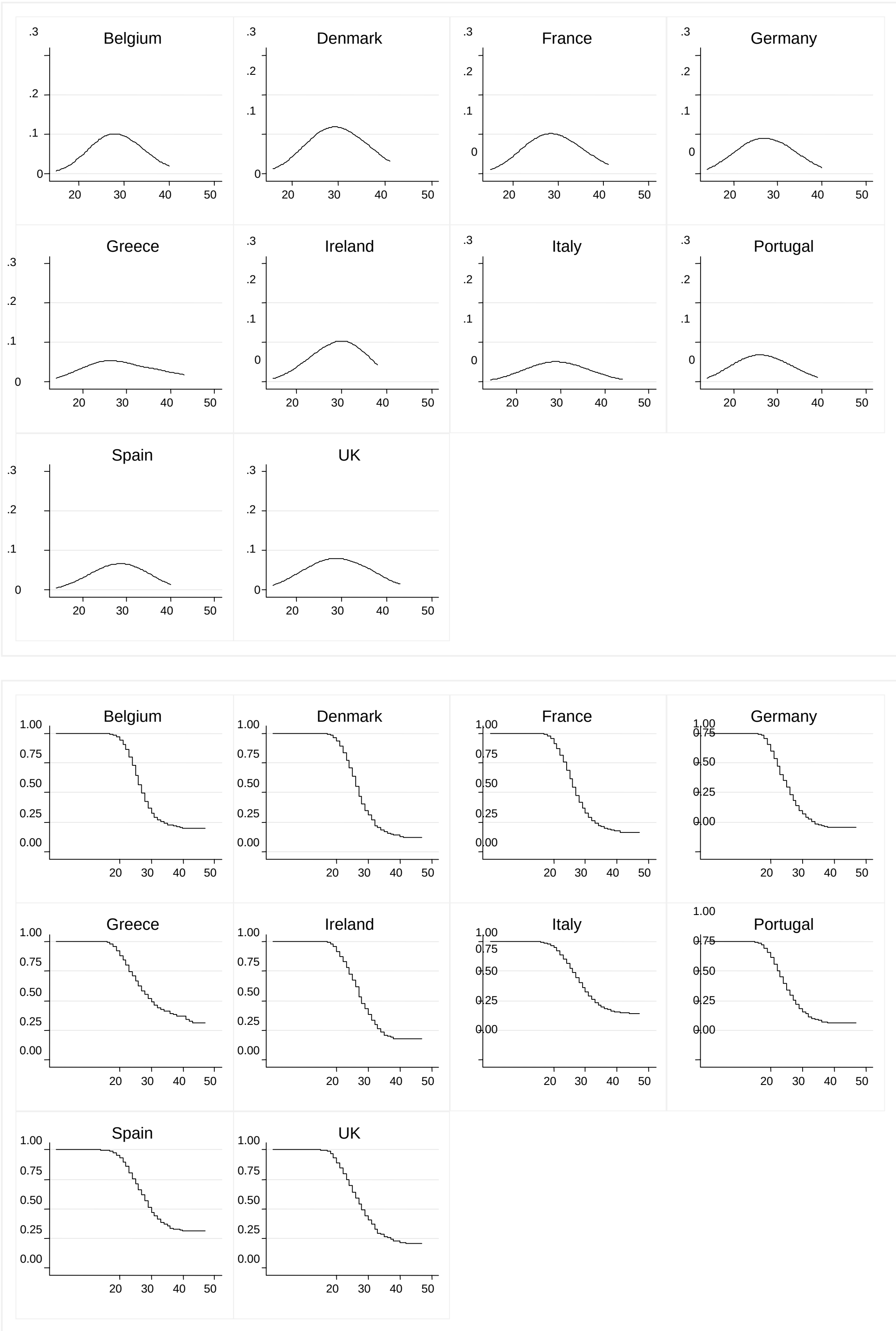
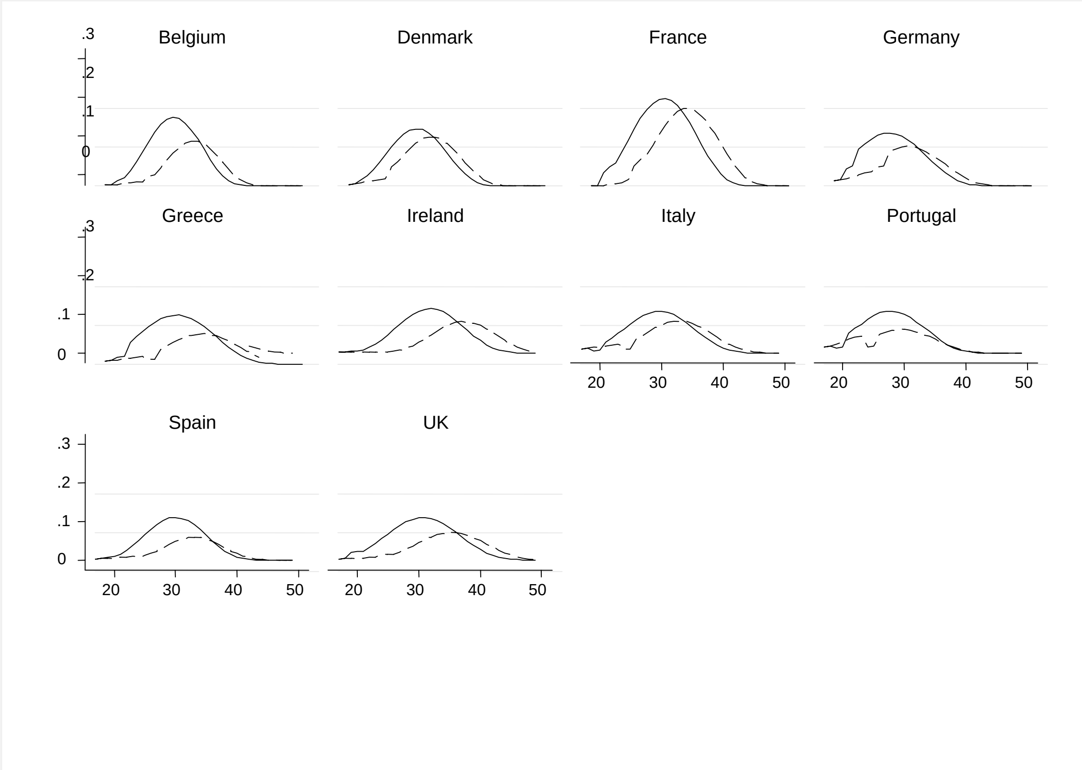
APPENDIX

Figure 1. First childbirth hazard and survival function estimated nonparametrically[9]



[9] 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)[10]



Tab. 1 Average age at 1st child, at 1st job and at the highest level of education completion

Country	Age at 1st child	Age at 1st job	Never worked	Age at hle
Belgium	25.55	20.69	0.11	20.53
Denmark	25.61	18.82	0.05	23.04
France	24.97	19.66	0.09	18.54
Germany	23.75	20.14	0.03	21.96
Greece	23.89	21.25	0.31	17.71
Ireland	24.95	18.11	0.12	18.70
Italy	25.62	20.76	0.34	18.42
Portugal	23.59	19.02	0.21	16.87
Spain	25.05	19.24	0.21	18.81
UK	25.30	18.27	0.07	17.90

[10] 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.

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	-2981.629	1005.816	0.000
Denmark	-7.23	0.000**	0.384	0.000**	-0.007	0.000**	10856	-2114.771	596.785	0.000
France	-6.417	0.000**	0.355	0.000**	-0.006	0.000**	30252	-5647.084	1959.938	0.000
Germany	-4.21	0.000**	0.298	0.000**	-0.006	0.000**	29183	-6314.116	827.309	0.000
Greece	-1.821	0.000**	0.116	0.001**	-0.001	-0.056	26884	-4204.158	1003.641	0.000
Ireland	-1.198	0.042*	-0.065	-0.217	0.001	-0.470	20320	-3525.287	1075.87	0.000
Italy	-4.683	0.000**	0.201	0.000**	-0.004	0.000**	45415	-5843.344	1433.509	0.000
Portugal	-4.981	0.000**	0.3	0.000**	-0.006	0.000**	22229	-4025.044	831.654	0.000
Spain	-5.09	0.000**	0.241	0.000**	-0.005	0.000**	40409	-6097	1585.69	0.000
UK	-2.751	0.000**	0.102	0.003**	-0.002	0.004**	23703	-4678.128	725.215	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.