STUDIA DEMOGRAFICZNE
1–2(183–184) 2024

Original research papers

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DOI: 10.33119/SD.2023.1-2.2

Labour market flexibility, education and completed fertility in 14 European countries

Abstract

Factors determining fertility are multidimensional, ranging from individual characteristics such as educational attainment and broader childbearing intentions to the socio-economic context and institutional support provided by family policies in a given country. Another important determinant of fertility is individuals' professional career stability which can be measured by the number of job changes.

The aim of this study is to investigate the associations between the number of children of mothers with the existing employment history (excluding childless women), education and labour market characteristics, individual career histories and job satisfaction in 14 European countries: Austria, Belgium, Czechia, Denmark, France, Germany, Greece, Ireland, Italy, the Netherlands, Poland, Spain, Sweden and Switzerland. In particular, labour market characteristics and career histories examined in this study are analysed with a focus on individual and country-level flexibility of the labour market.

It was found that education is negatively associated with the number of children of mothers aged 47 or older in analysed countries. Fertility patterns of highly educated mothers appear to be influenced by different factors than those of less educated mothers. Professional career stability and job satisfaction seem to have a relatively strong influence on mothers with tertiary education – as indicated by instrumental variable zero-truncated Poisson models. Moreover, divergent results connected with the employment flexibility of the individual and country-level brought about twofold conclusion: relatively less stable professional careers of mothers are negatively associated with the number of children: frequent job changes were generally undesirable by the middle-aged and older cohorts of mothers. Nevertheless, a more rigid labour market is also associated with lower completed fertility. Country-level flexibility, as measured by bargaining union coverage, appears to reflect labour market rigidity – hindering mothers' ability to return to the labour market post-childbirth – rather than stability that protects mothers from dropping out of the labour market. Finally, including economic status in the analysis indicated that job changes between the ages of 25 and 29 have the strongest impact on the completed fertility of 1910s–1960s female birth cohorts.

Keywords: education, fertility, labour market flexibility, Poisson regression, SHARE

Introduction

In the literature, the cross-country variation in fertility is attributed, for example, to micro-level education and employment conditions. One of the main ways in which economists explain the childbearing intentions is through labour supply and education (Becker, 1965). According to the microeconomic model of fertility, higher education aligns with increased earnings capacity and entails greater opportunity costs of having children, thus decreasing fertility. However, if children are considered normal goods (granting their parents a certain psychic income), higher earnings translate into more financial resources available for child-rearing, potentially increasing fertility. The first of these two divergent effects (substitution effect) was observed in particular for women (Butz, Ward, 1979). Importantly, in the 1970s, micro-level employment conditions were found to be increasingly important in influencing both quantum and tempo of fertility over the years (Karaman, Örsal, Goldstein, 2010).

At the same time, notions such as household economic situation or economic uncertainty are highly debated in socio-demographic studies. It should be noted that the literature on economic uncertainty as a factor explaining decreasing fertility has grown in popularity in recent years (Kreyenfeld, Andersson, Pailhé, 2012).

Education and labour market situation tend to interact with each other in their influence on fertility decisions. At the same time, job stability seems particularly relevant for higher fertility, especially for more educated women (Adsera, 2011). Importantly, employment stability was found to be a more important predictor of fertility for women than merely the state of being unemployed (Alderotti, Vignoli, Baccini, Matysiak, 2021).

One of the notions used in this study and related to employment stability is labour market flexibility, defined in different ways depending on the study (Kaplanis, Monastiriotis, 2012). Four different approaches to studying labour market flexibility require a brief discussion. These approaches include internal flexibility (covering flexible working hours arrangements), external (dealing with permanent and casual employment status), numerical (concerning shiftwork) and functional (job rotation and skill integration) (Wickramasinghe, Wickramasinghe, De Silva, Chanrasekara, 2019). These approaches reflect two axes across which the notion of labour market flexibility is often analysed in the literature: internal-external and numerical-functional axes. External numerical flexibility, which is investigated in this study, reflects the adjustability of labour in-take, for example, through temporary contracts. In certain theoretical frameworks, labour market flexibility is divided simply into time-of-work flexibility, wage flexibility, supply flexibility and employment flexibility, the latter being analogous to the external numerical flexibility. More specifically, employment flexibility refers to the division of careers into the ones consisting of frequent job changes and more stable ones. In the literature, employment flexibility is described as the most important measure of the broader labour market flexibility (Majewska, Samol, 2016).

Due to the dynamic nature of employment, studies on the association between labour market characteristics, education and fertility commonly use event history analysis, adopting a life-course perspective. Such studies are, however, usually limited to individual highly developed countries (Huinink, Kohli, 2014). Moreover, childlessness levels among the 1970 birth cohorts of women are not unusually high in comparison to the 1900 cohorts even though the tempo-adjusted fertility decreased (Sobotka, 2017). Hence, studies covering multiple countries and focusing on the number of children among mothers with at least one child seem vital. In addition, studies dealing with employment instability and fertility most commonly use unemployment or the fact of having temporary contracts as indicators measuring employment instability. This study will incorporate the number of job changes as a measure of career instability.

The aim of this work is to investigate the association between the number of children of mothers with the existing employment history (excluding childless women), education and labour market characteristics, individual career histories and job satisfaction in 14 European countries: Austria, Belgium, Czechia, Denmark, France, Germany, Greece, Ireland, Italy, Netherlands, Poland, Spain, Sweden and Switzerland.

In particular, labour market characteristics and career histories considered in this study are analysed with the focus on individual and country-level flexibility of the labour market. The role of economic status, interaction between education and career stability as well as differences between geographical regions will also be examined.

Literature review

According to the theory of planned behaviour, social norms, perception of behavioural control as well as behavioural attitudes and intentions constitute the most significant predictors of actual behaviour (Ajzen, 1991). Thus, considering specifically the impact of individual intentions on childbearing decisions, these decisions seem strongly dependent on individual characteristics and life choices of women, especially those concerning their professional careers.

One crucial factor determining the individual magnitude and quality of employability skills is represented by the level of educational attainment obtained by individuals. Indeed, educational attainment seems to be highly associated with fertility. Meta-studies on education and fertility show that as cohort fertility generally decreases in the twentieth century all around the world, no matter the educational attainment level of women, the fertility gap between low and highly educated women persists, as shown for instance by the statistics from such developed countries as Belgium, Germany, Japan, Norway and Sweden (Skirbekk, 2008). Depending on the country, one could also observe the shift in the relationship between fertility and education started in the 17th and 18th century in Western Europe. In the past, higher social status and education level were associated with a higher number of children in certain countries and it was only after the early modern historical period that higher education began to be negatively correlated with fertility. The switch in the number of children of people from higher social classes (compared to people from the lower classes) was demonstrated, for instance, in the analysis of the fertility patterns of different social strata of the French city of Rouen between the late 17th century and the French Revolution outbreak (Bardet, 1983).

A decline in fertility observed in Europe since the 1960s was accompanied, for example, by a rapid increase in the share of adults with tertiary education. The negative effect of enrolment in education on fertility has been observed in numerous studies. Such a negative influence was also found to be especially strong for women, tending to postpone motherhood in order to attain a desired education – thereby ensuring higher earnings and more stable labour market positions. Furthermore, women active in the labour market are more likely to have greater opportunity costs of childbearing if they are highly educated. At the same time, having children is also associated with increased likelihood of exiting the educational system. Delayed motherhood leaves less time for higher parity births. There is also evidence that, in recent years, the association between fertility and education has become less negative in highly developed countries (Wood, Neels, Kil, 2014).

Moreover, regional discrepancies are noticeable – for instance, Central and Eastern European states are characterised by more negative effects of education on the number of second and third births, compared to Northern, Western and Southern Europe (Wood et al., 2014). Likewise, Central and Eastern Europe differs from Western Europe by relatively low levels of childlessness and age at first birth of cohorts born after the World War II. In fact, common trends concerning the direction of changes in childlessness or higher-order births may be characteristic of the whole nation clusters or welfare state types. Nevertheless, the extent of fertility changes in different educational groups is rather time and country-specific or even more local (Brzozowska, 2015).

Studies have also compared fertility of couples depending on their union type and education level. Using hurdle zero-truncated Poisson models for three European countries, Osiewalska (2017) investigated the reproductive behaviour of unions that are homogamous, hypergamous or hypogamous in terms of their education level – unions characterised by the same level of education shared by both partners (homogamy), higher education level of men (female hypergamy) and higher education of women (female hypogamy). Hypergamy was found to be insignificant for higher fertility in all three countries. Moreover, the fertility patterns of hypergamous couples was similar to that of homogamous unions with medium-level education (Osiewalska, 2017).

The effects of education on fertility are often studied using the instrumental variable econometric methods that allow researchers to deal with the possible endogeneity bias related to the unobserved factors determining educational attainment level. One of such analyses, of Fort, Schneeweis and Winter-Ebmer (2016), based on the SHARE database (for Western Continental Europe) and ELSA database (for England), compared the results from the simple ordinary least squared (OLS) and Poisson econometric models with the two-stage least squares (2SLS) models, using an instrumental variable reflecting the mandatory schooling years required by law. The results indicated that controlling for endogeneity might mitigate or even neutralise the negative association between education and fertility in Austria, Denmark, Netherlands and Italy, but not in England (Fort et al., 2016). Other papers list common third factors influencing both fertility and education. The most important ones include household's financial situation, urbanisation, and life expectancy (Becker, Cinnirella, Woesmann, 2009).

One of the major topics in demographic research focuses on the association between fertility and micro-level employment conditions. The post-war employment conditions and labour market in general experienced significant changes. In particular, there was an expansion in the labour force participation of women (particularly married women and mothers). Importantly, the share of female-headed households also increased. Higher earnings and associated opportunity costs raised female employment rates, encouraged marital breakups and reduced the desire to have large families (Becker, 1991). Hence, the relationship between fertility and labour market conditions is nothing but complex: Recent labour market developments have influenced both employment opportunities and work-family arrangements (reshaping the traditional male breadwinner model), affecting fertility in divergent ways.

On the one hand, greater job opportunities, associated with higher earnings, are crucial to provide for a larger family. It was found that woman's financial contribution to the household budget may stabilise a marriage. Moreover, if both partners provide for a family, the couple is better prepared for a unexpected job loss of one of them (Oppenheimer, 1997). On the other hand, according to preference theory, the group of work-centred women (constituting up to 20% of female population), who realise themselves through work rather than family, may consider work and childbearing as competitive. Greater job opportunities may also partially influence fertility desires of adaptive women (around 60% of female population), who balance their work at home and in the workplace. The remaining lifestyle group of home-centred women is likely less affected by the expansion of job opportunities (Hakim, 2000). Mothers' lifestyle preferences, intentions, personal values and attitudes towards motherhood, jobs and family roles are central in shaping mothers' reproductive behaviour. A preference for work-life balance over career success seems to grow in popularity as women choose part-time jobs more often (Hakim, 2004). Women may also aim to postpone childbirth to first establish a stable position in the labour market so as not to jeopardise their career prospects (Kotowska Jóźwiak, Matysiak, Baranowska, 2008). Studies of the complex relationship between employment and fertility show that having children exerts a strong negative impact on mothers' employment. Likewise, employment reduces childbearing risk. Nevertheless, the latter effect is negligible for childless women (Matysiak, 2009).

Direction of the association between micro-level employment conditions and fertility may depend on the particular job characteristics. One of such characteristics, studied more often since the Great Recession is the stability of women professional careers. Employment instability and uncertainty have been increasingly prevalent in Europe since the 1973 oil crisis. There are different measures used to assess the extent of employment instability, including, primarily, time-limited contracts and unemployment. For instance, one article based on the EU-SILC data showed that holding a permanent (rather than temporary) job contract increases the likelihood of having first child for both Italian men and women (Vignoli, Drefahl, De Santis, 2012). Another study connecting joblessness with fertility indicated that persistent joblessness negatively affects woman's fertility intentions and is related to lower number of children or even higher likelihood of remaining childless (Busetta, Mednola, Vignoli, 2019). Employment instability forces people to postpone leaving their parental home and, as a consequence, postpones the decision to have children, leaving less time for higher-order births. The decision to have second or higher-order child may be also delayed or hindered by employment instability or uncertainty (as the decision to have higher-order children is more strongly determined by family's budget than by emotional motivations). This was shown by Alderotti et al. (2021) in the meta-analysis based on 49 selected articles concerning European countries and dealing with the association between employment instability and fertility.

In addition to the literature focusing on unemployment or temporary contracts as measures of employment instability, there are also studies investigating effects of job displacements on fertility. It turns out that fertility decline related to unexpected job displacements does not result from income losses generated by unemployment, nor from the length and incidence of individual unemployment spells. The losses in fertility seem to be rather associated with the job displacement *per se*, indicating a burden imposed by employment uncertainty on fertility. Furthermore, job losses may result in as much as 5 to 10 percent reductions in average fertility. It was shown, for instance, in the analyses based on the Austrian Social Security Database (Del Bono, Weber, Winter-Ebmer, 2012, 2015). It is worth noting that job displacements or, more generally, career interruptions may be also related to educational level as those with low educational attainment level tend to have more precarious and temporary jobs.

Moving on, the country-specific protective government legislation which may facilitate or impede the return to work after parental leaves may also indirectly influence fertility in an ambiguous way. For instance, the rigid employment protection legislation or excessive bargaining union coverage, on the one hand, aim at increasing the employment stability. On the other hand, they also increase wages, pushing them above the market-clearing level, leading to higher unemployment (exceeding the natural unemployment level and becoming compulsory for people with relatively lower skill level) (Snower, Lindbeck, 2001). The more rigid labour market may discourage employers from reducing employment in recessions and hiring new employees in the periods of economic growth. This may lead to an increased difficulty of mothers who want to return to the job market and thus promote the male breadwinner household model – indirectly impacting fertility, not necessarily in a positive way – as indicated by

Zhou and Kan (2019). Moreover, in the literature, the general flexibility of the labour market is usually described as a factor that allows women to combine professional careers and motherhood more easily (Doepke, Hannusch, Kindermann, Tertilt, 2022).

Education and labour market situation tend to interact in their influence on fertility decisions. For instance, fertility is found to respond more strongly to higher unemployment in a group of women with a relatively low level of education compared to highly educated women. Less educated women are also found to have a higher likelihood of a second child birth when they choose to remain inactive instead of remaining unemployed. Meanwhile, more permanent and secure jobs seem to be particularly relevant to more educated women in terms of higher fertility (Adsera, 2011).

Apart from individual fertility intentions and underlying factors, one can also consider the role of social norms and their impact on fertility desires. A subjective perception of social norms is one of the most important predictors of behaviour mentioned in the theory of planned behaviour (Ajzen, 1991). Moreover, the individual is more likely to maintain or adapt a certain behaviour when it is perceived by them as relatively easy (Murray, Häubl, 2007). Specifically for childbearing and child-rearing, the more societies ensure that caretaking is perceived as easy to undertake, the higher the fertility intentions (Matsuo, Matthijs, 2010). Thus, childbearing decisions of women are also highly dependent on the sociocultural and economic background, incentivising motherhood – via social norms and further institutional support for such norms (for example, family policy tools encouraging or reshaping the traditional male-breadwinner and female-caretaker household model). Family policy legislation of a country constitutes a significant factor determining fertility rate. For instance, the introduction of paid parental leave may impact women's labour market behaviour in several, often diverging, ways, making the total effect on their childbearing decisions ambiguous. On the one hand, longer leaves may facilitate the return to work of mothers who prefer to stay at home for a longer period of time, and would have abandoned the labour force had their leave been shorter (Klerman, Leibowitz, 1999). On the other hand, parental leaves give mothers an incentive for pre-birth work in order to remain eligible for benefits. Childbirth paid leave also increases the value of the time spent by women executing childrearing responsibilities at home and thus increases the reservation wage above which women are willing to return to the labour force (Boeri, van Ours, 2013).

A number of studies revolve around the effects of family and employee benefits on fertility and employment patterns. In their comparison of parental leaves in three Nordic countries, Rønsen and Sundström (2002) show that family policies play a prominent role in shaping mothers' after-birth labour-market behaviour. Results indicated that despite the general effectiveness in enhancing women's ability to remain in the labour market throughout the childbearing years, family policies have to be designed carefully: too generous parental leaves may prolong women's career breaks, whereas too long leaves may strengthen the unequal division of domestic work. On the other hand, too weak pre-birth incentives can make it difficult for women to reconcile their childbearing and occupational responsibilities. Improperly designed policies can be detrimental to both women's careers and earnings potential. Nevertheless, it was also found that women who are entitled to a paid leave have a much higher overall employment entry rate (Rønsen, Sundström, 2002).

Even after a quick glance at the literature concerning the association between family policies and childbearing decisions, one can come across analyses using the SHARE database. In one of such studies, Brugiavini, Pasini and Trevisan (2013), investigated the impact of maternity leave policies on female labour market attachment (time spent at home after childbirth). In their analysis, authors rearranged data of the SHARELIFE database into a birth panel. In this panel, each maternity episode constituted a single observation. The information about the country and age of mothers, as well as the year of such episodes, made it possible to account for the specific family policies that could influence these mothers' decisions to stay at home for a longer period of time and the potential outcomes of such decisions. The analysis indicated that more generous paid maternity leave reduces the number of weeks spent outside of the labour market, even when it results in a prolonged period of time spent at home after childbirth (Brugiavini et al., 2013). These findings support the more general aforementioned conclusion drawn earlier by Klerman and Leibowitz (1999) – the one concerning the alleviation of financial burden of mothers who want to re-enter the labour market after a slightly prolonged time spent at home post-childbirth.

Method and data

The two-stage instrumental variable (IV) regression, estimated using the generalised method of moments (GMM), was selected to address the complicated relationship between completed fertility, education and labour market characteristics. The data used for this study were obtained from the SHARELIFE database (SHARE Wave 3) concerning the retrospective data on women aged 47 or older. Some questions from which the variables were calculated originate from other SHARE rounds, especially the data about job satisfaction obtained from the SHARE Wave 8. The underlying model used in the second stage of the regression was the zero-truncated Poisson model in which clustered standard errors were calculated at the country level. The estimation was weighted using the cross-sectional weights from the third round of SHARE.

The choice of the zero-truncated Poisson count data model was dictated by the nature of the dependent variable: the number of children of each woman (excluding childless respondents). This variable could only take non-negative integer values and the value of zero was impossible. The total number of observations used in the analysis was necessarily relatively low due to the usage of data derived from different SHARE rounds and, in one of the models, amounted to only 2537 observations. This was unavoidable, given the fact that over half of SHARE respondents participated in one interview only (Fort et al., 2016).

The main model included 3385 female respondents who had at least one child. They were born between the 1910s and the 1960s, with the majority born in the 1950s (55.9%) and the 1940s (37.1%). The proportions of respondents with 1, 2, 3, or 4+ children were 18.2%, 47.9%, 22.7%, and 11.1%, respectively. Only 120 respondents (3.6%) had never been married. A relatively large share of respondents resided in Sweden (12.1%), Denmark (11.3%), France (10.1%), and the Netherlands (8.5%), whereas comparatively few were citizens of Ireland (2.3%), Austria (2.8%), Poland (4.5%), and Spain (5.0%).

It should be noted that the decision to exclude childless women from the analysis was made at an early stage. An additional Poisson-logit hurdle regression model, including childless women (presented in Table 1), indicated that different factors influence, to varying degrees, the likelihood of having at least one child (versus remaining childless), and having more children overall. In particular, unlike subjective financial wellbeing, education is more strongly negatively associated with having higher-order children than with having at least one child. Moreover, higher expenditure on family benefits has a significant positive effect on the number of children only for families with at least one child. To narrow the focus of the analysis, this paper primarily considers mothers with at least one child, excluding childless women. This narrower scope also allows for a closer examination of the impact of delayed motherhood on the number of children.

Table 1. Poisson-logit hurdle model (including childless women)

	Number of children (dependent variable)
Logit component	
Age	–0.016*** (0.004)
Ever married	1.338*** (0.201)

	Number of children (dependent variable)
Subjective financial wellbeing	0.042* (0.024)
Number of job changes	–0.047*** (0.019)
Education	–0.126*** (0.035)
Family benefits expenditure (percentage of the GDP)	0.080 (0.094)
bargaining coverage	–0.022*** (0.008)
Constant	3.137*** (0.008)
Poisson component	
Age	0.012*** (0.003)
Ever married	–0.193* (0.110)
Subjective financial wellbeing	0.007 (0.007)
Number of job changes	–0.050*** (0.013)
Education	–0.135*** (0.023)
Family benefits expenditure (percentage of the GDP)	0.130* (0.073)
Bargaining coverage	–0.013** (0.006)
Constant	0.615 (0.432)
Observations	6948

Note: Robust standard errors adjusted for 14 clusters (countries) in parentheses, * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.
Source: Author's own work.

The fertility-education association explored in this paper may be affected by endogeneity bias, especially in the form of simultaneity bias – decisions concerning the number of children and investment in education are likely simultaneously determined by common factors. Ignoring the suspected simultaneity may result in a correlation between regressors and the error term leading to biased and inconsistent estimates. Becker, Cinnirella and Woesmann (2009) mentioned three common factors influencing fertility and education at the same time: life expectancy, urbanisation rate and income. Urbanisation may partly affect the choice between more education and more fertility. More urban environments tend to be more individualistic (value orientation towards large and extended families is lower), are arguably less suited to childrearing and less family-friendly (for example due to proximity of natural and rural areas) and involve higher opportunity costs for parents. Furthermore, household's financial situation may strongly determine both childbearing decisions and decisions concerning investment in education. As regards life expectancy, one of economic theoretical perspectives dealing with home-based decision-making (belonging to a set of theories referred to as the New Home Economics) suggests that households' decisions on how much to invest in education of children may be influenced by the expected length of productive life. Life expectancy is also associated with such correlates as the level of social development, quality of healthcare and possibilities to adopt a healthy lifestyle.

In the context of instrumental variable analysis, it should be noted that the three aforementioned common factors of education and fertility should be part of the error term so as not to bias the education parameter. Nevertheless, inclusion of variables differentiated at the country level (country of respondent birth) might potentially make some of these macro-level common factors valid as instruments in the fertility-education nexus (robust tests for weak instruments are particularly useful to check their relevance).

Considering that the data used in this analysis concern middle-aged and older women, gender inequality may constitute a potential factor influencing fertility mainly through its impact on education (or, more specifically, access to education). It is, however, also at risk of being too much of a common factor of fertility and education. According to the literature, gender equity may explain the choice of either more education or more fertility (Galor, Weil, 1996). Moreover, women engaging in relatively more housework tend to have lower fertility intentions, especially when their job satisfaction is low (Mills, 2008). Nevertheless, in certain European countries, such as Czechia, policy legislation promoting the male breadwinner – female caregiver model (in the form of the so-called explicit familialism) turned out to be beneficial for the total period fertility rate and for the relatively quick transition to older fertility patterns (Szelewa, 2007), compared to the neighbouring Central European countries. Therefore, the association of gender inequality with female education seems less ambiguous (relative to its association with fertility), indicating that gender inequality might potentially constitute a valid instrument.

Apart from the simultaneity problem, it should be noted that since measurement of potential factors affecting fertility may be subject to imperfections, the omitted variable bias may also arise, making error terms correlated with fertility.

All independent variables apart from the education level were treated as exogenous. The exogenous variables used in different models calculated in this study included

1. instruments:
   - age – respondent's age;
   - urbanisation rate – the percentage of urban population in respondent's country of birth in 1990;
   - female life expectancy at birth – life expectancy at birth of women calculated for the period between 1985 and 1990 obtained from the United Nations database;
   - GDP per capita – gross domestic product (GDP) per capita based on the purchasing power parity in the respondent's country of birth in 1985 obtained from the International Monetary Fund database;
   - gender inequality – Gender Inequality Index of 1990 in the respondent's country of birth; a measure reflecting gender disparities in health, empowerment and the labour market, obtained from the United Nations Development Programme database;
2. and the remaining variables:
   - number of job changes – number of job changes of a respondent added to reflect the stability of mothers' professional career;
   - age at first childbirth – age at which mothers had their first child;
   - job satisfaction – categorical variable reflecting the level of job satisfaction of respondents who could state that they strongly disagree, disagree, agree or strongly agree that they are satisfied with their current job;
   - family benefits expenditure (percentage of the GDP) – the expenditure on family policies in relation to their GDP in 1980s or 1990s (depending on data availability) in the respondent's country of birth; this variable was obtained from OECD Stat and Eurostat databases;
   - bargaining coverage – bargaining union coverage in the labour market of the respondent's country of birth; it was calculated as the number of employees covered by valid collective (wage) bargaining agreements divided by the number of all wage and salary earners, and obtained from the Institutional Characteristics of Trade Unions, Wage Setting, State Intervention and Social Pacts (ICTWSS) database; this variable was added to measure the general employee's bargaining power and, indirectly, the lack of employment flexibility in the labour market;
   - ever married – binary variable equal to unity if a woman was married at least once in her life and zero otherwise;
   - subjective financial wellbeing – reflecting the subjectively assessed difficulty of the respondent (considering the total monthly income) to make ends meet financially; respondents who stated that they were able to make ends meet

with a great difficulty, difficulty, fairly easily and easily were assigned the respective values of 1, 2, 3 or 4.

The variable measuring the number of job changes needs a more detailed description: each episode of permanent employment or series of temporary jobs (exceeding a total of 6 months) was counted as a separate job. Individuals were considered to change their jobs if they changed an employer or only changed their position with the same employer but indicated in the questionnaire that this change should be classified as a different job. In addition, maternity leaves were not included as job interruptions. Jobs were also considered separate if the respondent performed more than one job at the same time but the period of performing one job did not entirely coincide with the period of performing another one.

The endogenous variable, education, assigns one of six levels of the highest educational attainment to each respondent in accordance with the 1997 International Standard Classification of Education (ISCED'97). In the case of mothers who stated that they are still in education or that they did not accomplish the primary education (the ISCED level indicated in the SHARE database was presented as 'None'), the value of this variable was set to zero.

Econometric Models and Their Properties

In order to explore the association between completed fertility, education and labour market characteristics, nine different instrumental variable (IV) zero-truncated Poisson regressions were computed.

An important and initial aspect of the analytical part of the study was the selection of instruments. A range of variables differentiated at the individual and country-level were tested as potential instruments. These correlates originated from SHARE, Eurostat, the Global Political Demography Database, ICTWSS, International Monetary Fund database, OECDStat and Survey of Adult Skills (PIAAC). The set of potential instruments was reduced based on conceptual reasons and tests for overidentification, underidentification and instrument weakness. Testing for instrument validity was complicated due to the type of estimation and inclusion of both individual and country-level regressors. Zero-truncated Poisson regression made it possible to calculate tests for overidentifying restrictions but the remaining tests were performed using auxiliary linear IV regressions. To account for heteroskedasticity across countries, clustered standard errors were computed, which necessitated the use of robust tests for underidentification (Kleibergen-Paap rk Wald F statistic). Inference robust to weak instruments was also performed using the Conditional Likelihood Ratio (CLR) and the Anderson-Rubin tests.

In the first auxiliary regression (allowing to consider the problem of weak instruments more easily), the education variable was treated as continuous, whereas in the second estimation – as categorical. Values of the categorical version of the education variable were aggregated into three broader ranges of educational attainment levels: first – comprising pre-primary, primary and lower secondary education (ISCED levels 0, 1 and 2); second – comprising upper and post-secondary education (ISCED levels 3 and 4); and third – comprising tertiary education (ISCED levels 5 and 6).[1]

It should be noted that all variables suspected of being prime common factors of fertility and education (GDP per capita, urbanisation and life expectancy) were excluded from the analysis – partly because of this suspicion and partly due to their relative weakness. For instance, clustering of standard errors at the country level made GDP per capita insignificant and increased the standard error of the urbanisation variable so much that it was also on the verge of statistical significance.[2] The Kleibergen-Paap Lm statistic of underidentification indicated that country-level female life expectancy at birth should be excluded from the analysis as well. The exclusion of the latter resulted in a slight decrease in the significance level of almost all regressors including the instrumented variable (education level).

Eventually, respondent's age and country-level gender inequality were the only two variables selected as instruments for the first set of models in this study (Tables 2–6). The F-Statistic of Wald test (of excluded instruments) in the first stage of the auxiliary regression (Table 2) was equal to 94.11, exceeding a sufficient level of correlation with the endogenous variable – indicating that the problem of weak instruments is not present (Staiger, Stock, 1997). Moreover, all instrumental variables in the first-stage regression were statistically significant with the respective p-values in significance tests lower than 0.05. Overidentification tests were not needed because the instrumented variable was categorical and could take three different values.

Furthermore, variance inflation factor (VIF) statistics calculated for both stages of estimation indicated that the problem of collinearity is not present in the model. Moreover, all independent variables in the second-stage estimation (except one category of the job satisfaction variable) were statistically significant at the level of

[1] These three categories of educational attainment were represented respectively by 32, 38 and 30% of respondents included in the econometric analysis.

[2] Nevertheless, the validity of the country-level urbanisation rate as a strong instrument in the individual-level education-fertility nexus was visible in additional models having less variables but more observations. Moreover, urbanisation was correlated with education much more strongly than with fertility.

significance of 0.1. The values of coefficients, their significance, standard errors, VIF statistics as well as R-squared and F-statistic obtained in the auxiliary first-stage estimation are presented in Table 2.

The values of coefficients, their significance and standard errors obtained in the second-stage estimation of the auxiliary model are presented in Table 3. Similar results connected with the second-stage equation of the main model are presented in Table 4.

Table 2. The first-stage estimation (education treated as the continuous variable)

	Education (endogenous variable)	Variance Inflation Factor
Age	–0.016*** (0.001)	1.03
Gender inequality	–1.987** (0.765)	1.03
Constant	3.512*** (0.291)	
Observations	3,385	
R-squared	0.03	
F(2, 13)	94.11	

Note: Robust standard errors adjusted for 14 clusters (countries) in parentheses, * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.
Source: Author's own work.

Table 3. The second-stage estimation (education treated as the continuous variable)

	Number of children (dependent variable)
Education	–0.502* (0.265)
Number of job changes	–0.036*** (0.008)
Age at first childbirth	–0.022** (0.011)
Ever married	0.313* (0.181)
Job satisfaction (ref.: strongly disagree)	
Disagree	0.227* (0.132)
Agree	0.341** (0.174)
Strongly agree	0.349*** (0.134)
Family benefits expenditure (percentage of the GDP)	0.178*** (0.047)
Bargaining coverage	–0.011*** (0.003)

	Number of children (dependent variable)
Constant	1.658*** (0.411)
Observations	3385

Note: Robust standard errors adjusted for 14 clusters (countries) in parentheses, * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.
Instruments: age, gender inequality.

Source: Author's own work.

Table 4. The second-stage estimation (education treated as the categorical variable)

	Number of children (dependent variable)
Education (ref.: pre-primary, primary or lower secondary)	
Upper and post-secondary	–0.882* (0.534)
Tertiary	–0.927* (0.528)
Number of job changes	–0.029*** (0.012)
Age at first childbirth	–0.021* (0.013)
Ever married	0.282* (0.156)
Job satisfaction (ref.: strongly disagree)	
Disagree	0.273 (0.173)
Agree	0.315* (0.167)
Strongly agree	0.360*** (0.147)
Family benefits expenditure (percentage of the GDP)	0.181*** (0.052)
Bargaining coverage	–0.011*** (0.003)
Constant	1.270*** (0.426)
Observations	3385

Note: Robust standard errors adjusted for 14 clusters (countries) in parentheses, * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.
Instruments: age, gender inequality.
Underidentification test in the auxiliary linear regression (Kleibergen-Paap rk LM): 4.692, p-value: 0.0958.
Conditional Likelihood Ratio test of the joint significance of the endogenous regressors in the auxiliary linear regression: 2.82, p-value: 0.099; Anderson-Rubin test of the joint significance of the endogenous regressors in the auxiliary linear regression: 3.03, p-value: 0.220.

Source: Author's own work.

For the purpose of considering the timing of job changes, the variable reflecting the number of job changes was disaggregated into three different age groups: below 25 years, from 25 to 29 years, and 30 years and above. The results are presented in Table 5.

Table 5. The second-stage estimation (education treated as the categorical variable) with job changes in different age groups

	Number of children (dependent variable)
Education (ref.: pre-primary, primary or lower secondary)	
Upper and post-secondary	–0.951* (0.583)
Tertiary	–0.892* (0.525)
Number of job changes (up to 24 years)	–0.082** (0.038)
Number of job changes (25–29 years)	–0.049 (0.039)
Number of job changes (30 years and more)	0.009 (0.029)
Age at first childbirth	–0.022* (0.013)
Ever married	0.287** (0.151)
Job satisfaction (ref.: strongly disagree)	
Disagree	0.281 (0.179)
Agree	0.316* (0.176)
Strongly agree	0.365*** (0.153)
Family benefits expenditure (percentage of the GDP)	0.184*** (0.055)
Bargaining coverage	–0.011*** (0.003)
Constant	1.402*** (0.423)
Observations	3380

Note: Robust standard errors adjusted for 14 clusters (countries) in parentheses, * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.
Instruments: age, gender inequality.
Underidentification test in the auxiliary linear regression (Kleibergen-Paap rk LM): 4.543, p-value: 0.1032.
Kleibergen-Paap rk Wald F statistic testing weak identification in the auxiliary linear regression: 34.520 (higher than typical Stock-Yogo weak ID test critical values).
Conditional Likelihood Ratio test of the joint significance of the endogenous regressors in the auxiliary linear regression: 2.93, p-value: 0.088; Anderson-Rubin test of the joint significance of the endogenous regressors in the auxiliary linear regression: 3.10, p-value: 0.212.

Source: Author's own work.

Table 6. The second-stage estimation excluding number of job changes

	Number of children (dependent variable)
Education (ref.: pre-primary, primary or lower secondary)	
Upper and post-secondary	–0.861* (0.522)
Tertiary	–1.127* (0.600)
Age at first childbirth	–0.018 (0.012)
Ever married	0.285* (0.157)
Job satisfaction (ref.: strongly disagree)	
Disagree	0.260 (0.177)
Agree	0.315* (0.166)
Strongly agree	0.361*** (0.118)
Family benefits expenditure (percentage of the GDP)	0.174*** (0.052)
Bargaining coverage	–0.010*** (0.003)
Constant	1.117*** (0.472)
Observations	3385

Note: Robust standard errors adjusted for 14 clusters (countries) in parentheses, * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.
Instruments: age, gender inequality.
Underidentification test in the auxiliary linear regression (Kleibergen-Paap rk LM): 4.674, p-value: 0.097.
Kleibergen-Paap rk Wald F statistic testing weak identification in the auxiliary linear regression: 44.608 (higher than typical Stock-Yogo weak ID test critical values).
Conditional Likelihood Ratio test of the joint significance of the endogenous regressors in the auxiliary linear regression: 2.79, p-value: 0.115; Anderson-Rubin test of the joint significance of the endogenous regressors in the auxiliary linear regression: 3.22, p-value: 0.200.

Source: Author's own work.

In order to check the interaction between the education level and the number of job changes, two supplementary Poisson regressions were calculated. The first one differed from the main estimation (presented in Table 4) in that the number of job changes was excluded from it. Similarly, the education variable was excluded from the second one. Both supplementary regressions are juxtaposed in Tables 6 and 7, respectively.

It was also verified whether the estimates from Table 4 would change substantially if education (the instrumented variable) were replaced by a more general economic status. The economic status variable was obtained by combining education and subjective financial wellbeing, where the latter variable reflected the subjectively assessed difficulty of the respondent in making ends meet financially. Linear combinations of these two variables were created using the principal component analysis. The first principal component, explaining 58% of the total variance of education and subjective financial wellbeing, was used in the zero-truncated IV Poisson regression from Table 8. Apart from attempting to find the specification that reflects reality more accurately (without increasing the model's complexity), the idea to combine these variables is also partially related to the fact that the polychoric correlation between them was found to be moderate (equal to 0.178) but relatively high compared to parametric or nonparametric correlations calculated for pairs of other variables in the analysis. The verification of underidentification and weak instrument problem indicated that the most optimal set of instruments for economic status should include age, gender inequality and female life expectancy at birth.[3] These instruments were used in the analysis.

Table 7. The second-stage estimation excluding education

	Number of children (dependent variable)
Number of job changes	-0.032^{***} (0.005)
Age at first childbirth	-0.034^{***} (0.008)
Ever married	0.255 (0.170)
Job satisfaction (ref.: strongly disagree)	
Disagree	0.212^{*} (0.113)
Agree	0.191 (0.160)
Strongly agree	0.203^{*} (0.113)
Family benefits expenditure (percentage of the GDP)	0.088^{**} (0.041)
Bargaining coverage	-0.009^{***} (0.002)
Constant	1.219^{***} (0.170)
Observations	3435

Note: Robust standard errors adjusted for 14 clusters (countries) in parentheses, * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.
Source: Authors' own work.

3 Female life expectancy was assumed to pose less risk of being a common factor in the endogeneity problem between fertility and economic status – compared to the endogeneity problem between fertility and education.

Table 8. The second-stage estimation including economic status (combining education and subjective financial wellbeing)

	Number of children (dependent variable)
Economic status	-0.211^{***} (0.054)
Number of job changes	-0.027^{***} (0.004)
Age at first childbirth	-0.035^{***} (0.003)
Ever married	0.240 (0.165)
Job satisfaction (ref.: strongly disagree)	
Disagree	0.177 (0.137)
Agree	0.219 (0.175)
Strongly agree	0.333^{***} (0.127)
Family benefits expenditure (percentage of the GDP)	0.244^{***} (0.051)
Bargaining coverage	-0.016^{***} (0.003)
Constant	1.356^{***} (0.310)
Observations	2517

Note: Robust standard errors adjusted for 14 clusters (countries) in parentheses, * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.
Instruments: age, gender inequality, female life expectancy at birth.
Hansen's J statistic testing for overidentification: 2.353, p-value: 0.308.
Underidentification test in the auxiliary linear regression (Kleibergen-Paap rk LM): 7.624, p-value: 0.0545.
Kleibergen-Paap rk Wald F statistic testing weak identification in the auxiliary linear regression: 19.666 (lower than one Stock-Yogo weak ID test critical value – for 10% maximal IV size).
Conditional Likelihood Ratio test of the joint significance of the endogenous regressors in the auxiliary linear regression: 0.35, p-value: 0.567; Anderson-Rubin test of the joint significance of the endogenous regressors in the auxiliary linear regression: 7.94, p-value: 0.047.

Source: Author's own work.

Since three instruments were used in the analysis, it was necessary to test for overidentification. The J test for overidentifying restrictions indicated that, at the 5-percent level of significance, there was no evidence against the null hypothesis of instrument validity – the Sargan-Hansen J statistic was 2.353, yielding a p-value of 0.308. After adding female life expectancy as an instrument, a potential issue regarding the statistical significance of the economic status variable emerged – as indicated by the Conditional Likelihood Ratio test but not by the Anderson-Rubin test (a less powerful one). However, this issue did not appear in the models accounting for cross-country differences.

Similarly to Table 5, Table 9 contains the results of the previous model complemented with separate variables concerning the timing of job changes.

Table 9. The second-stage estimation including economic status and job changes in different age groups

	Number of children (dependent variable)
Economic status	–0.195*** (0.058)
Number of job changes (up to 24 years)	–0.052*** (0.014)
Number of job changes (25–29 years)	–0.062*** (0.015)
Number of job changes (30 years and more)	–0.006 (0.009)
Age at first childbirth	–0.036*** (0.003)
Ever married	0.249 (0.161)
Job satisfaction (ref.: strongly disagree)	
Disagree	0.184 (0.139)
Agree	0.210 (0.174)
Strongly agree	0.326*** (0.126)
Family benefits expenditure (percentage of the GDP)	0.238*** (0.053)
Bargaining coverage	–0.015*** (0.003)
Constant	1.416*** (0.310)
Observations	2513

Note: Robust standard errors adjusted for 14 clusters (countries) in parentheses, * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.
Instruments: age, gender inequality, female life expectancy at birth.
Hansen's J statistic testing for overidentification: 2.353, p-value: 0.308.
Underidentification test in the auxiliary linear regression (Kleibergen-Paap rk LM): 7.794, p-value: 0.050.
Kleibergen-Paap rk Wald F statistic testing weak identification in the auxiliary linear regression: 17.745 (lower than one Stock-Yogo weak ID test critical value – for 10% maximal IV size).
Conditional Likelihood Ratio test of the joint significance of the endogenous regressors in the auxiliary linear regression: 0.48, p-value: 0.492; Anderson-Rubin test of the joint significance of the endogenous regressors in the auxiliary linear regression: 7.72, p-value: 0.052.

Source: Author's own work.

The last two models were computed in order to obtain results concerning differences in fertility between countries. Due to a relatively low number of observations, fourteen countries were grouped into three regime types according to Esping-Andersen's (1990) welfare state typology. More specifically, countries were grouped into conservative corporate (Continental), social democratic (Scandinavian) and rudimentary (South European). Ireland was assigned to Continental and not the British (liberal) type of welfare state because of the low number of respondents from this country and the lack of other countries with the liberal regime. It should be noted that the specification of the last models was slightly different from previous regressions – two country-level variables, family benefits expenditure and bargaining coverage, were excluded from the analysis. Moreover, age at first childbirth, gender inequality and female life expectancy at birth were selected as instruments after performing and comparing the properties of several alternative estimations with different specifications. It should be noted that, at the level of significance of 0.05, an overidentification problem could be present, as indicated by the Sargan-Hansen J statistic. Nevertheless, at lower significance levels (for example 0.025), overidentification was not evident. The first stage estimation raised slight doubts about the significance of life expectancy – the variable ensuring the statistical significance of instrumented economic status (indicating a potential problem of weak identification). However, importantly, inclusion of life expectancy did not influence other variables, especially the number of job changes. Poisson regression comparing different welfare state regimes is presented in Table 10. Its extension taking into account the timing of job changes is displayed in Table 11.

Table 10. The second-stage estimation including welfare state types

	Number of children (dependent variable)
Economic status	–0.249*** (0.104)
Number of job changes	–0.031*** (0.005)
Age	0.017*** (0.006)
Ever married	0.278 (0.206)
Job satisfaction (ref.: strongly disagree)	
Disagree	0.318** (0.158)
Agree	0.279 (0.187)
Strongly agree	0.288** (0.139)

cont. Table 10

	Number of children (dependent variable)
Welfare state type (ref.: conservative corporate / continental)	
Social democratic / Scandinavian	0.205*** (0.067)
Rudimentary / South European	–0.326** (0.138)
Constant	–1.520*** (0.522)
Observations	2517

Note: Robust standard errors adjusted for 14 clusters (countries) in parentheses, * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.
Instruments: age at first childbirth, gender inequality, female life expectancy at birth.
Hansen's J statistic testing for overidentification: 7.414, p-value: 0.025.
Underidentification test in the auxiliary linear regression (Kleibergen-Paap rk LM): 7.715, p-value: 0.052.
Kleibergen-Paap rk Wald F statistic testing weak identification in the auxiliary linear regression: 31.712 (higher than typical Stock-Yogo weak ID test critical values).
Conditional Likelihood Ratio test of the joint significance of the endogenous regressors in the auxiliary linear regression: 31.63, p-value: 0.000; Anderson-Rubin test of the joint significance of the endogenous regressors in the auxiliary linear regression: 34.72, p-value: 0.000.

Source: Author's own work.

Table 11. The second-stage estimation including welfare state types and job changes in different age groups

	Number of children (dependent variable)
Economic status	–0.231*** (0.105)
Number of job changes (up to 24 years)	–0.064** (0.027)
Number of job changes (25–29 years)	–0.082*** (0.024)
Number of job changes (30 years and more)	0.0005 (0.010)
Age	0.018*** (0.006)
Ever married	0.283 (0.199)
Job satisfaction (ref.: strongly disagree)	
Disagree	0.323** (0.163)
Agree	0.279 (0.187)
Strongly agree	0.283** (0.138)
Welfare state type (ref.: conservative corporate / continental)	
Social democratic / Scandinavian	0.196*** (0.066)

	Number of children (dependent variable)
Rudimentary / South European	−0.313** (0.132)
Constant	−1.506*** (0.524)
Observations	2513

Note: Robust standard errors adjusted for 14 clusters (countries) in parentheses, * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.
Instruments: age at first childbirth, gender inequality, female life expectancy at birth.
Hansen's J statistic testing for overidentification: 7.553, p-value: 0.023.
Underidentification test in the auxiliary linear regression (Kleibergen-Paap rk LM): 7.866, p-value: 0.049.
Kleibergen-Paap rk Wald F statistic testing weak identification in the auxiliary linear regression: 31.616 (higher than typical Stock-Yogo weak ID test critical values).
Conditional Likelihood Ratio test of the joint significance of the endogenous regressors in the auxiliary linear regression: 31.18, p-value: 0.000; Anderson-Rubin test of the joint significance of the endogenous regressors in the auxiliary linear regression: 34.15, p-value: 0.000.

Source: Author's own work.

Results

The coefficient of the education variable in the main estimation showed that education is negatively associated with the number of children of mothers aged 47 or older in 14 analysed countries: the comparison of coefficients for two categories of education indicates that, compared with the least educated group, tertiary education reduces mothers' expected number of children by more than does secondary (and post-secondary) education, *ceteris paribus*.

The reduction of statistical significance following the elimination of instruments reflecting country-level urbanisation and life expectancy was stronger for mothers with the secondary education – indicating that the bias related to common factors was smaller for mothers with the tertiary education. This indicates that common factors affecting female fertility (not accounting for childless women) differ depending on the level of education. These factors are also potentially less typical (not related to common factors cited in the literature) as the level of education increases.

Moreover, some additional findings are connected with GDP per capita. This variable was considered a weak instrument and was excluded from analysis at an early stage. It is noteworthy, however, that the possible bias associated with the inclusion of GDP per capita affected primarily the coefficient for the tertiary education. Estimates for education instrumented by GDP per capita even pointed towards the possible U-shaped association, or at least a flattening of the association between education and fertility (the difference in the average number of children between mothers with tertiary education and mothers from the least educated group was lower than the difference between mothers with secondary and post-secondary education, and the least educated group). It is worth noting that similar results regarding the flattening of the association between education and fertility were obtained in the model from Table 5 – taking into account the timing of job changes.

In addition, if one compares models from Table 4 and Table 6 (the latter excluding the variable reflecting the number of job changes), it becomes evident that the discrepancy between values of coefficients of secondary and tertiary education is significantly larger in the model that does not account for job changes. Importantly, the addition of the number of job changes influenced mainly the coefficient for tertiary education. This might indicate that, in the context of increased fertility, career stability is more relevant to highly educated mothers than to those with lower education attainment levels. All these findings suggest that fertility of mothers with the highest education level is affected by different factors than fertility of less educated mothers. Identifying and accounting for these factors might result in a flattening of the negative fertility-education association for higher education attainment levels. Professional career stability seems to constitute one of such factors. Conversely, as indicated in Table 7, differences in education do not influence the relationship between professional career stability and fertility.

The finding concerning the relative importance of career stability for more educated women in terms of higher fertility is similar to that of Adsera (2011). However, job stability was measured in a more unconventional way – reflecting the number of job changes and not unemployment or temporary employment.

Furthermore, the fact that the negative association between education and completed fertility remains significant (and not mitigated after considering the endogeneity problem) distinguishes the obtained results from those highlighted by Fort et al. (2016) in their Poisson regression based on the SHARE database (with a lower set of countries analysed and using different instruments).

The results connected with employment flexibility at the individual and country level (reflected by the number of job changes and bargaining coverage respectively) are divergent. The relatively less stable professional careers of mothers, characterised by the frequent job changes, are negatively associated with the number of children.[4] Nevertheless, the more rigid labour market also goes in line with lower fertility. The explanation of such results may be that the frequent changes of jobs were rather unwanted by the middle-aged and elderly cohorts (the results might be different

[4] One should bear in mind that the negative value of coefficient of the number of job changes would likely be smaller if one was to additionally account for partner characteristics – in line with Alderotti et al. (2021, p. 886).

in the case of younger cohorts). Thus, the number of job changes should be regarded as a measure of career instability. More frequent job changes do not reflect ease of returning to the labour market after a career break following a childbirth. The same could not be said about bargaining coverage which measures rather a labour market rigidity (resulting in the difficulty of mothers returning to the labour market after a break related to childbirth) than stability (protecting mothers from dropping out of the labour market).

Both results indicate that family legislation should address the parental work-life balance challenges in a comprehensive way, rather than a purely pro-natalist one (focused mainly on material aid). In particular, policymakers should not disregard the importance of ensuring the career stability of mothers, bearing in mind that such tools as paid maternity leaves should be designed carefully, as indicated by Rønsen and Sundström (2002). One tool that was not discussed in this study but may be significant in facilitating the combination of career and childbirth is the labour market policies – helping individuals flow from unemployment or inactivity to employment. Especially active labour market policies (public and intervention works, loans or subsidies granted to entrepreneurs or the unemployed, training programmes, as well as job search assistance) are found to be effective in counteracting compulsory unemployment (Majewska, Samol 2016).

Accounting for the timing of job changes in Tables 5, 9 and 11 shows that the impact of career instability on completed fertility of 1910s–1960s female birth cohorts is the strongest when job changes take place before the age of thirty. Results from Tables 9 and 11 (robust to modifications in the set of instruments) show that job changes carried out between the ages of 25 and 29 have the strongest impact on completed fertility in fourteen analysed European countries.

A one-year postponement of the age at first birth is associated with an incidence rate ratio (IRR) of 0.965–0.979 for the total number of children per mother, *ceteris paribus*. A coefficient of each threshold of the job satisfaction categorical variable indicates that job satisfaction is intuitively positively associated with fertility. Nevertheless, exclusion of education in the model from Table 7 made strong job satisfaction a less positive and less significant factor of working mothers' fertility. This indicates that differences in education levels were relevant for the higher fertility, especially for mothers who were most satisfied with their work. Higher job satisfaction also seems more important for higher educated women (hypothetical exclusion of the job satisfaction variable changed mainly the coefficient for tertiary education). Moreover, one could draw the following conclusion related to the positive sign of the family benefits expenditure variable. The well-funded comprehensive family policy model seems to be generally more effective than the less popular non-interventionist model, reducing citizens' dependency on state benefits and promoting private welfare benefits (represented, for instance, by the British government in the 1980s and 1990s as mentioned by Chiu, Wong and Yip (2008)).

The findings related to instruments used in the analysis raise doubts about inclusion of such variables as GDP per capita, urbanisation and life expectancy in the analysis of fertility-education association, even if these three correlates are differentiated at the country level. Considering GDP per capita, urbanisation and life expectancy at the country level may mitigate their problem as common factors in the fertility-education nexus (because they are correlated more strongly with individual-level education than completed fertility) but these three factors cause a problem of weak identification. Nevertheless, it was found that Gender Inequality Index is of a high importance in explaining completed fertility through its effect on different education attainment levels. Validity of gender inequality as an instrument indicates that the association between a country's gender equality and fertility is mainly mediated through effects of this variable exerted on education (the model containing gender inequality as a non-instrumental independent variable was less suited to the empirical data and led to an underidentification problem). Using gender inequality as an instrumental variable is, however, not without its caveats – it may influence completed fertility through other channels than only education attainment of mothers, for example through labour market opportunities.

Furthermore, as indicated by the coefficient of determination in Table 2, only 3% of variation in education is preserved in the model from Table 3 (slightly more in the model from Table 4, taking into account the categorical nature of the education variable). The comparison of Tables 1 and 3 shows that the impact of education on fertility in the Poisson-logit model was likely underestimated. Bearing in mind that the direction of the bias in the association between education and fertility depends on the type of omitted confounder (Kamhöfer, Westphal, 2017), the estimate for education in the IV model from Table 4 may be slightly downwardly biased itself: Poisson-logit does not directly account for family's wealth or openness to experience, which may downwardly bias the estimate. The omission of the household wealth was addressed in the model from Table 8.

Coefficients and standard errors in regression including economic status (Table 8) did not change significantly. Thus, bearing in mind that the number of observations decreased to only 2517 from 3385 due to the inclusion of the subjective financial wellbeing variable, the specification of the model containing economic status instead of education may be better suited to empirical data. Future analyses may seek to replace the relatively weak instrument of female life expectancy at birth and account for confounders not discussed in this study, especially psychological traits from the "gv_big5" SHARE module. Accounting for different territorial units smaller than countries seems promising as well. The only obstacle is that the number of observation drops rapidly due to the inclusion of variables from different SHARE rounds.

Finally, the model in Table 10 indicated differences in fertility between European welfare states of three different types. Social democratic welfare states turned out to be the most encouraging for bigger families. Conservative corporate welfare states were slightly less favourable in terms of higher fertility of mothers (with a history of being active in the labour market). Meanwhile, life in rudimentary welfare states was associated with the least favourable environment for having more children.

Conclusions

The analysis indicates that education is negatively associated with the number of children of mothers aged 47 or older in fourteen analysed countries. Fertility of mothers with the highest education level is influenced by different factors (which may be atypical vis-à-vis the literature on common macroeconomic determinants of fertility and education) than fertility of less educated mothers. Professional career stability and job satisfaction seem to constitute such correlates – influencing completed fertility of highly educated mothers more strongly than that of mothers with lower education. At the same time, the impact of career instability on completed fertility seems to be the strongest when job changes occur between the ages of 25 and 29.

The divergent results connected with the employment flexibility at the individual and country level lead to twofold conclusions: The relatively less stable professional careers of mothers are negatively associated with the number of children. Nevertheless, the more rigid labour market also goes in line with lower fertility. The explanation of such results may be that the frequent changes of jobs were rather unwanted by the middle-aged and elderly cohorts. Meanwhile, bargaining union coverage measures rather a labour market rigidity (resulting in the difficulty of mothers' returning to the labour market after a break related to childbirth) than stability (protecting mothers from dropping out of the labour market).

Models from Tables 6 and 7 (investigating the interaction between education and the number of job changes) showed that even though professional career stability is relatively important for completed fertility of highly educated mothers, differences in education do not influence the relationship between professional career stability and fertility.

The higher age at first birth is associated with the lower number of children. Job satisfaction is intuitively positively associated with fertility. Given the results concerning the family benefits expenditure, a well-funded family policy seems to be generally more effective than (for instance) non-interventionist family policies.

Conclusions related to the instruments used in the analysis indicate that gender inequality may constitute a valid country-level instrument in the education–childbearing endogeneity problem. One major limitation of this instrument – and of the others, particularly age (birth cohort) – concerns the exclusion restriction: these variables may affect the number of children not only through education, but also through other channels such as women's labour-market opportunities, access to contraception, norms surrounding ideal family size or marriage timing. Due to SHARE data limitations, it was also not possible to include such potential instruments as the month-of-birth variable which would measure individuals' biological age within the academic or calendar year. Moreover, both age and the Gender Inequality Index may function as proxies for broader societal gender-related norms that vary across cohorts and countries, potentially influencing fertility mainly through their impact on educational attainment rather than exerting a direct effect on fertility (thus acting more as instruments than confounders) – at least for cohorts born between the 1910s and the 1960s. Furthermore, economic status, combining education and subjective financial wellbeing, might lead to a specification which more accurately models the analysed associations. The model containing economic status was, however, affected by a drop in the number of observations related to the addition of the subjective financial wellbeing variable. A different set of instrumental variables was required in the estimation of the model with economic status compared to the model with education.

Finally, welfare states in Northern Europe (Scandinavia) seem more encouraging for higher fertility of mothers than those in Southern Europe.

To sum up, factors determining completed fertility are multidimensional and related to both individual characteristics and life-course choices, as well as the country-specific context. The general negative association between education and fertility observed since the 17th century was found to be supported in the 21st century SHARE data for middle-aged and elderly mothers (with the history of being active in the labour market) in the 14 analysed European countries, despite diverging results obtained in studies applying similar methods of count variable regressions. Moreover, a carefully designed and well-financed, but not necessarily pro-natalist, family policy seems to be conducive to an increased fertility rate. Nevertheless, it should be pointed out that family policy and labour market policies should be aligned in promoting the stable (and satisfying) professional careers of mothers – such careers were found to be relevant primarily to highly educated mothers. Future studies can focus on the impact of labour market policies on fertility while accounting for the individual frequency of job changes. Future analyses based on SHARE may account for such confounders as psychological traits as well as different territorial units smaller than countries. Considering the results connected with the bargaining coverage variable, ensuring a more flexible labour market should be favourable to mothers' professional careers.

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