Four Essays on Demographic Economics

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"People modify their behavior not only in response to changing external (market) conditions, but also in response to what they have done and to what has happened to them in the past"

Richard H. Day

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Chapter 5

Abstract

The objective of this thesis is to probe four demographic events from an economic perspective. Specifically, we employ the Easterlin relative income hypothesis to provide some explanation on movements regarding marriage, non-marital fertility, and fertility postponement. Easterlin claims that young adults decide on several issues based on their relative affluence. By the latter, Easterlin implies that young adults compare their current (potential) economic condition to the one experienced during their childhood. Thus, the higher the affluence enjoyed in childhood, the higher their demands in young adulthood. This is, in short, the Easterlin relative income hypothesis. We retrieve data from IPUMS-CPS for the period 1981-2016 across the US for white non-Hispanics (due to data availability) to provide some evidence on the latter. Results corroborate the hypothesis. Relative income of young adults found statistically significant for all the three aforementioned demographic events. In particular, relative income is related negatively to non-marital fertility and women's fertility postponement, but positively to marriage rates. In addition, we find that relative income behaves better than the absolute one (in size and statistical significance aspects) with respect to marriage and premarital births. Next, we examine the fertility rebound. The latter took place in late 90s early 00's especially in the most developed countries. Conventional wisdom tests for the rebound with respect to HDI (Human Development Index) and GDP (Gross Domestic Product) per capita. We differentiate by testing the rebound in terms of labour productivity and female labour force participation, except for GPD per capita. Our analysis focuses on the turning points of two developed OECD income country groups (high and low income) that spans the period 1970-2016. The aim is to find out the factor (among labour productivity, female labour force participation, and GDP per capita) for which the turning points of the rebound (if confirmed) between the two groups, are closer, or coincide (statistically insignificant). Results show that differences in turning points are closer for labour productivity rather than GDP per capita or female labour force participation. Some thoughts on why the latter might hold are provided at the first part of the paper, based on demography, economic theory, and recent evidence. We conclude that labour productivity

arises as the most important economic factor for the onset of the fertility rebound.

CHAPTER 1

Introduction

This dissertation consists of four papers. Three of them refer to the Easterlin relative income hypothesis on three contemporary major demographic events: marriage, non-marital fertility, and fertility postponement. The last paper is related indirectly to Easterlin. It has been inspired by Easterlin's relativity perspective on demography, and suggests, to the best of our knowledge, a new explication for the fertility rebound phenomenon.

The first paper discusses the issue of marriage within the Easterlin relative income hypothesis frame. In this paper, we follow the critique of Macunovich for the incorrect implementation of the Easterlin hypothesis; especially in marriage literature. Macunovich claims that there have been two studies that test for the Easterlin hypothesis in its original outset. One study provides evidence in favour of the hypothesis whilst the other against. Therefore, result is ambiguous. We add on the relevant literature by presenting some new evidence on the Easterlin relative income hypothesis in the context of marriage. The hypothesis has been examined for the US at a period that spans 1981-2016 by employing panel data analysis and causality tests. The results confirm the hypothesis stated: a drop in the relative income of young men leads to marriage declines. In addition, we compare the effect of the relative income on marriage with that of the absolute one. We find that relative income exerts a greater impact on marriage than the latter.

The second paper is very closely related to the first one. Here, we aim to answer to the question of whether young adults' relative income – in the sense of Easterlin – coincides with the increase of the number of children born out-of-wedlock of young women. This is the first paper, to our knowledge, that seeks directly to correlate these two variables. The hypothesis is confirmed. Thus, as young adults feel less affluent compared to the level of their material aspirations (and before they adjust their aspirations to lower levels as a consequence of their fail to achieve them), the odds to a born outside marriage increases. As already mentioned, this is very closely related to the marriage-relative income concept but with a major difference: we incorporate marriage rates as a control variable to verify that relative income still impacts on births out-of-wedlock independent of it. One could reasonably argue that the statistically significant correlation found between non-marital fertility and relative income is the outcome of the reduced number of marriages: marriage attenuates and probabilities of birth out-of-wedlock increase. Hence, the examined relationship could be thought as a spurious one, driven by marriage. However, holding marriage rates constant (control variable), relative income remains statistically significant on non-marital fertility. This implies that the examined relationship is not spurious rather relative income exerts an effect on non-marital fertility through other channels too. We do not indicate the channels that might drive the relationship found. This is left for a future investigation.

The third paper focuses on the fertility postponement and the female education in the US. The latter (female education) was not in our intention to be examined in the first place. In this study, the primary idea had been the linkage of relative income to later childbearing. However, there had been two reasons to include female education in this paper.

First, we were very skeptic about the confined data availability of the IPUMS-CPS on the mean age at first birth. The results obtained for mean age at first birth were in line to the hypothesis stated. But, the low number of observations presented in the latter case reduces its reliability. Thus, we were seeking an additional way to strengthen these results.

Second, literature has so far shown that the most significant factor on women's fertility postponement is the increasing rate of their educational attainment (Ní Bhrolcháin and Beaujouan, 2012; Neels et al., 2017). Hence, we thought of some arguments on why relative income of young men would affect female education and we tested for it. Panel data analysis and Granger non-causality tests confirmed our hypothesis. Thus, we have incorporated female education into the paper in order to distinguish between direct and indirect impact of relative income on fertility postponement (by indirect we imply through female education). That means we have implicitly employed female education as an instrument to the fertility postponement due to data scarcity on the latter.

Accordingly, this paper examines two hypotheses which both refer to the fertility postponement. In the first step we examine the hypothesis directly: “The relative income decline of young men coincides with shifts to later childbearing (direct impact on the fertility postponement).” (p.66). In the second step we test the hypothesis indirectly through female education: “The relative income decline of young men coincides with increases to female educational attainment” (p.67). The contribution of this paper lies into those two aspects of social demography.

Before we move into the last paper, we have to clarify an important issue with respect to the previous ones. We have employed micro-data in aggregation form by state, year, and age. We should note that the Easterlin hypothesis has been conducted both within macro and micro framework (see Macunovich, 1998: 98). In both cases the hypothesis has been either rejected or confirmed. None is *a priori* false or correct. We adopt the macro approach for one important reason. Easterlin instruments aspirations by the family income during childhood. He also claims over a positive correlation between family income in childhood and the formation of the material aspirations. Hence, the higher the family income enjoyed in childhood, the higher the aspirations formed and transmitted to young adulthood. The latter is reasonable. However, what if someone has grown up into a very disadvantaged household? Accordingly, his aspirations should be lower than his peers who grew up in a more affluent environment. We argue that the latter might be true up to a point; below that point the family income-aspirations relationship could be reversed. That is, his relative deprivation (this time deprivation with respect to his peers) would induce him to acquire more material aspirations and thus be even more materialistic in young adulthood (Moschis 1987, 2007; Kasser et al., 1995; Moschis et al., 2009; Nguyen et al., 2009).

A recent paper of Whelan and Hingston (2018) confirm such thoughts. The latter show that adults who grew up as poor children usually place more importance on achieving the material norm than their wealthier peers. If the positive correlation was universal, we should not indicate such evidence. Complementary to that, we could also speculate on the opposite; on those who have grown up in highly affluent environments. These are individuals who usually get higher education, and maybe, higher education acts negatively on materialism. This is an assumption. We are not aware of any empirical studies that demonstrate it. If it holds, it would in turn imply that aspirations in childhood increase with family income up to a certain point; beyond that point the relationship might once again reverse (aspirations reduce). Thus, we claim that there might be two bounds: one on the upper and one on the lower strata. The positive relationship exists within these bounds and, perhaps, deviates outside. These thoughts led us to conduct the analysis in an aggregate way where such mechanisms probably do not take place and the hypothesis for a positive relationship seems more presumable. It has been left for a future work to test whether the hypothesis (upper-lower bound) stated in this short analysis holds or not.

In addition to the above, data constrain has also been an obstacle for a micro-oriented approach on the Easterlin hypothesis. We had checked the most well-known potential databases for this purpose: Panel Study of Income Dynamics (PSID), British Household Panel Survey (BHPS), Integrated Public Use Microdata Series (IPUMS-CPS), and European Social Survey (ESS), but without finding all the variables needed for an effective relative income hypothesis test within a micro frame.

The final paper of this dissertation is about the fertility rebound. This is a relative new topic in demography coming from the publication of Myrskylä et al. (2009). They show that the fertility rate reverses after some point of human advances (from low fertility rates to higher). The latter has been the onset for a debate in demography. Some researchers argue in favour of the rebound whereas others against. Both sides, however, address the rebound with respect to the Human Development Index (HDI) and the Gross Domestic Product (GDP) per capita. In this paper, apart from these two variables, we test for the rebound (the convex relationship between fertility and the other factors considered) in terms of labour productivity and female labour force participation. Inspired by Easterlin's relative income concept, we provide some arguments on why labour productivity is anticipated to be the most important factor for the onset of the rebound. Similarly to relative income, labour productivity is a ratio. It consists of the GDP divided by the sum of working hours. Thus, in line with Easterlin's thought (income per se is not an adequate index to infer on various demographic phenomena), we argue that GDP per capita per se may not be so strong predictor for the rebound effect as labour productivity. A combination of the GDP to the time consumed for its production and the ensuing leisure time might explicate it better. In other words, labour productivity depicts the relationship between income and working time in a macro perspective. We investigate this hypothesis by the comparison of the turning points between two income country groups (high and low) for GDP per capita, labour productivity, and female labour force participation. Indeed, results show that the distance between the turning points is minimized for labour productivity rather than GDP per capita or female labour force participation, in most of the cases examined. The latter implies that the fertility decline of a country reverses if it reaches a specific range of labour productivity rather than GDP per capita or women's labour participation. Labour productivity turns out as the decisive factor. Thus, regression analysis confirms that labour productivity could be thought as the main driver for the onset of the fertility rebound. Female labour force participation follows on later stages and strengthens it. This last paper concludes that labour productivity might answer on why we also see the fertility rebound in developing countries where the levels of female labour force participation are still low enough.

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CHAPTER 2

Why Young Adults Retreat from Marriage? An Easterlin Relative Income Approach

Abstract

Easterlin's relative income hypothesis refers to the current income of young adults compared to the level of material aspirations acquired during childhood. We examine how relative income affects marriage. A higher (lower) level of income or/and a lower (higher) level of material aspirations increases (decreases) relative income and consequently marriage rates. We employ panel data for the United States that spans the period 1981-2016 to test this hypothesis. Panel dynamic methods and causality tests are applied. We reveal that young men are more likely to wed if they feel affluent relative to the level of their material aspirations. Relative income emerges as a stronger predictor than the absolute one in all methods investigated.

Keywords: marriage, relative income, Easterlin hypothesis.

JEL Classification Codes: J11, J12, J19.

2.1 Introduction

United States (US) has experienced a steady decline in marriage rates over the last half century. According to Martin et al. (2014), one-third of young adults in their 20s will never marry. Marriage decline has also led to an increase in the fraction of children born out-of-wedlock and concerns over the family structure (Lerman, 1996). However, many young men still perceive marriage as one of the most important milestones of their life (Willoughby et al., 2015). This contradiction between data and young men's preference on marriage has led to a riddle called the "marriage paradox" (Willoughby and James, 2017). In Fig. 2.1 we see the percentage decline of marriage rates between 1990 and 2016 across states.

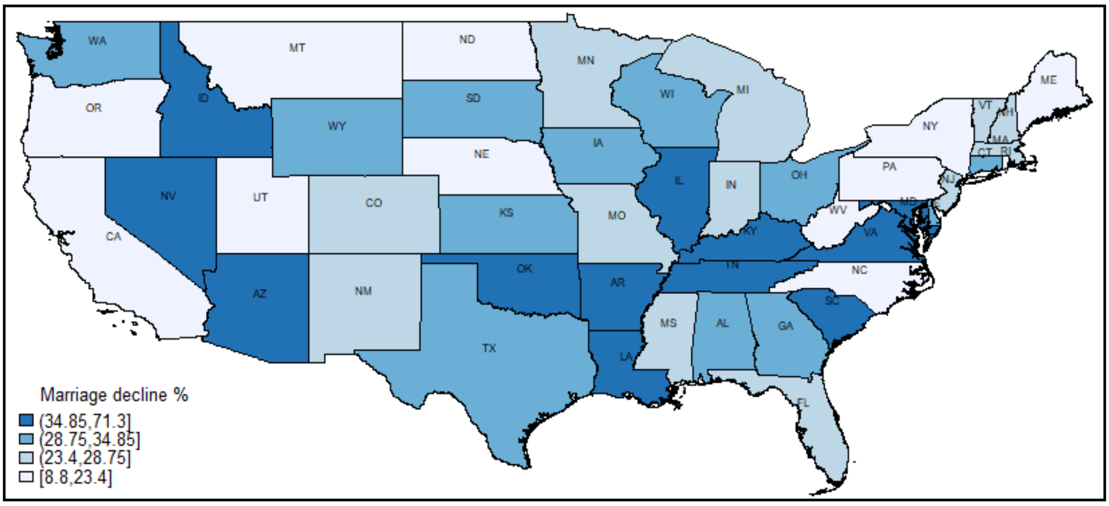

Fig. 2.1 The percentage decline of marriage rates across the US between 1990 and 2016. *Note*: Calculation of the percentage change has been conducted by the authors. *Source*: CDC/NCHS, National Vital Statistics System.

The attenuation of marriage is not an isolated phenomenon. It came along with the increase of cohabitation (Ishizuka, 2018), women's economic improvements (Becker, 1991; Cherlin, 1992), contraceptive access (McLanahan, 2004; Stevenson and Wolfers, 2007), and men's lower marriageability (Bridges and Boyd, 2016). The percentage of men aged between 18 and 24 years old living with spouse in the US reduced from 31.2 in 1967 to 5.1 in 2016 (Fig. 2.2). The percentage for women is 46.3 and 9.1, respectively. On the contrary, the percentage of young adults who live in partnership has increased (Fig. 2.2). The latter implies a trend to marriage delays (Fig. 2.3).

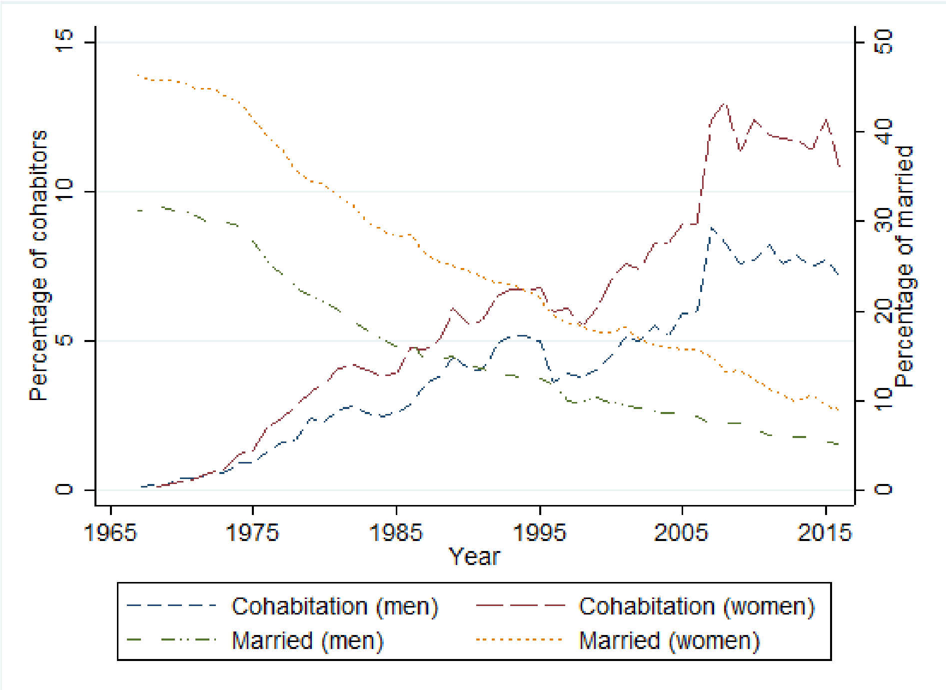

Fig. 2.2 Percentage of married and cohabiting young adults aged 18-24. *Source*: US Census Bureau, Current Population Survey, Annual Social and Economic Supplement.

The postponement in younger ages is also in line with Willoughby et al. (2015). Willoughby argues that marriage still remains important for young adults' life; more important than careers and leisure activities. Indeed, the decline is less prominent as we move to individuals who are over 50 (compare Fig. 2.3 and 2.4). Hence, one could claim that the drop we observe at younger ages may reflect economic considerations. Instead, the relatively smaller drop for the ones who are over 50 could mirror sources other than economic (ethical, political, religious, etc.). This evidence is aligned with Easterlin's hypothesis as we will explain in section 2.3.

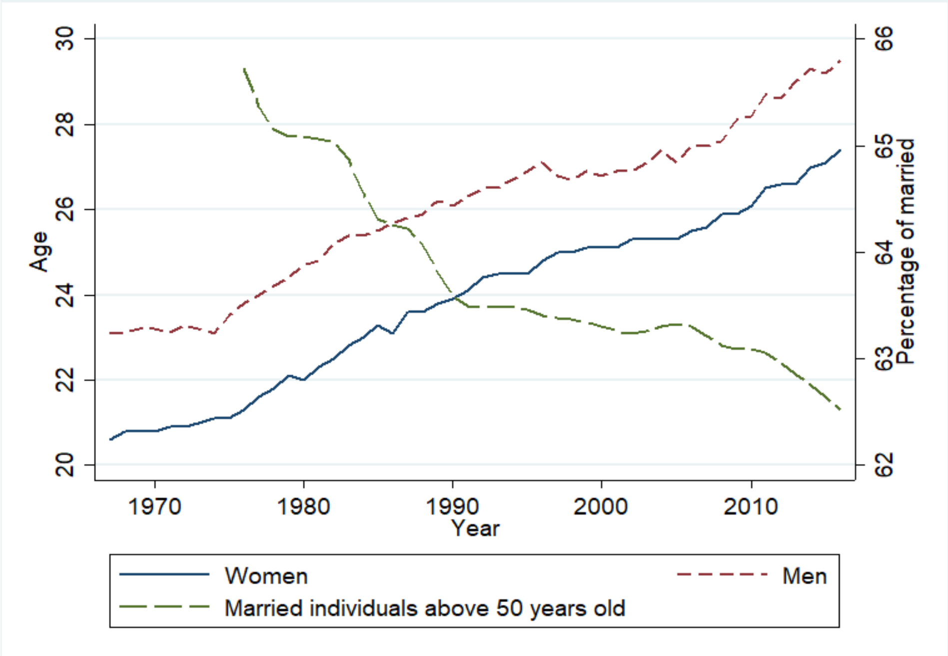

Fig. 2.3 Median age at first marriage for men and women, and the percentage of married individuals (white non-Hispanics) aged over 50 years old. *Source*: For the age at first marriage: US Census Bureau, Current Population Survey, March and Annual Social and Economic Supplements. For married individuals above 50: our own calculations employing data extracted from the IPUMS-CPS.

Yet, Easterlin's hypothesis has not always been applied in the marriage literature within its original context.[1] The purpose of this study is to fill this gap. We contribute to this strand of the literature by investigating why young adults in the US retreat from marriage. We reveal that the ratio computed by the earnings and the material aspirations[2] is an important factor for marriage. Thus, relative income might offer an explanation on young adults' retreat from marriage. That may also be a solution to the "marriage paradox".

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[1] For a criticism on the relativity measures that have been used to test Easterlin's hypothesis (not solely with respect to marriage), see Macunovich (1997; 1998a; 1998b).

[2] Hereafter we refer to material aspirations simply as aspirations.

2.2 Literature Review

The existing literature suggests a number of factors that contributed to the decline in marriage rates during the last decades in the US. Social scientists focus mainly on the importance of gender roles in marriage stability (Parsons, 1949; Pessin, 2018). Becker (1991) examines gender roles but he follows an economic approach. Inspired by the international trade theory, Becker argues that men and women will choose to marry only if the gains of marriage are higher than being single. He sees individuals as trade partners whose gains of marriage are maximized when each partner specializes in a specific domain: men in labour market and women in household work. Thus, women's socio-economic improvements through labour force participation, educational attainment, and increased earnings, attenuate their benefits from marriage and reduce its rates.

Becker's contention had a great appeal in economics and many attributed the retreat from marriage to women's emancipation (Ermisch, 1981; McLanahan and Casper, 1995; Sassler and Schoen, 1999; Sweeney, 2002). However, Oppenheimer (1988, 1997) questions the power of Becker's trade theory on marriage and provides evidence on the deterioration of young men's labour market position. Oppenheimer also notes that most aggregate cross-sectional studies find negative effects with respect to women impact, whereas individual-level longitudinal positive ones. Therefore, the impact of women's empowerment on marriage is rather mixed depending on the approach.

Another major theory on marriage concerns ideological issues. According to van de Kaa (1987), shifts in ethical, political, and religious beliefs are responsible for a large part of the marriage decline. Other explanations consider the effect of the "marriage squeeze" theory (gender ratio). This theory assumes that the existence of a greater number of women relative to the number of men would worsen chances for women to find a partner. Angrist (2002) tests the squeeze theory on marriage patterns in second generation immigrants and finds a positive effect of the gender ratio for women. Most recently, Bronson and Mazzocco (2018) suggest that changes in cohort size over time and across states explain about half of the variation in US marriage rates since the early twentieth century. Finally, technological changes on birth control such as pill contraceptives (Goldin and Katz, 2002) and abortion availability (Akerlof et al., 1996) affect the desire to marry.

2.3 Marriage in Easterlin's context

2.3.1 The Easterlin hypothesis

Easterlin (1966) claims that young adults decide on a number of aspects such as marriage and fertility based on their relative affluence. There is a consumption threshold for each generation (macro perspective) or for each adult (micro perspective) which has to be reached to proceed to marriage. The formation of the consumption threshold stems from childhood (Easterlin, 1987). Easterlin formulated his theory to interpret the post-World War II baby boom in the US. His theory posits cyclical changes in demographic and social behavior due to the fluctuations in birth rates. Thus, a large cohort will produce a small one due to unfavorable labour market conditions that it meets and a small cohort will obtain a large one, for the same reason. Easterlin developed two measures to test his hypothesis: the relative cohort size and the relative income (RY henceforth). The two concepts – despite their linkage – have been proved to be different.[3] A discrepancy between the two measures had already been observed since 1973 (Macunovich, 1998a). Therefore, Easterlin's theory could be seen from two perspectives. The first one relates to the relative cohort size and the second one to the RY.

This paper focuses on the second perspective. Easterlin discusses the role of aspirations for the decision process of young adults. Aspirations were proxied with the family income of their parents during childhood years (Easterlin, 1966). Easterlin argues that there is a positive correlation; the higher the family income

[3] For a discussion on the opposite direction of relative cohort size movements than that proposed by Easterlin, see Lutz et al. (2006).

in childhood, the higher the aspirations formed and transmitted to adulthood.[4] However, the main problem with the construction of RY is the formation of the denominator which is assumed to represent aspirations (Macunovich, 1997). Aspirations are primarily formed within household, but also affected by the influence of peers, neighborhood, school, television, and other sources (Richins, 2017). Hence, it is crucial to take into account the (average) income of all families in a specific area (e.g., state).

The conclusion of Easterlin's hypothesis is the following: young adults whose incomes are high enough compared to their desired consumption threshold[5] will feel freer and will proceed to marriage. Otherwise, they will choose to postpone – until they reach this threshold – or even forgone marriage.

2.3.2 The Easterlin hypothesis in the context of marriage

RY hypothesis has been widely tested in the fertility literature but it has received less attention with respect to marriage.[6] Macunovich (2011) contends that there are only two studies in the relevant literature which have tested directly the RY and marriage relationship: MacDonald and Rindfuss (1981), and Macunovich (2002).[7] She also observes that "there has been a wide diversity of measures which have been developed empirically and tested in the name of the 'Easterlin Hypothesis'" (Macunovich, 1997: 122). In a review paper Macunovich (1998a: 55) points out: "[…] an examination of the studies which provide least support might lead some to question whether they actually address the Easterlin

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4 The late work of Easterlin on the economics of happiness recognized that young adults may have the same level of aspirations independently of their family background (Easterlin, 2001).

5 For a clarification on the concept of childhood family income and its relation to consumption threshold, see Macunovich (1998a: 102).

6 See Pampel and Peters (1995: 180).

7 "There have been two studies that have tested this relationship between relative income and marriage directly – that is, using older family income: MacDonald and Rindfuss (1981) and Macunovich (2002), with the first finding no support for the theory, and the second finding strong support. Other studies have looked at the effect indirectly – using parental education and/or occupation as proxies for income." Macunovich (2011: 18).

hypothesis as he formulated it – but this could be said for several of the supportive studies as well! Sometimes because of data limitations, sometimes because of widely varying interpretations of the hypothesis, researchers have conducted studies which seem to bear little resemblance to the hypothesis".

MacDonald and Rindfuss (1981) investigated a sample of men who graduated from Wisconsin high schools in 1957 and found no support of the relative income hypothesis. They concluded that only their own income (absolute income) affects family formation. Schapiro (1988) finds his developed measure of RY to predict divorce but not marriage rates. Watson and McLanahan (2011) maintain that raising the reference group of men's incomes by ten percent reduces marriage by about two percent. In a similar study, Loughran (2002) finds that male wage inequality is correlated significantly to women's propensity to marry.

Another paper that tests the RY hypothesis more closely to its original concept is that of Macunovich (2011). Macunovich employed the parental family income of men and women currently married who were 0-5, 6-10, and 11-15 years out of school. Macunovich found a negative parental family income effect on marriage. However, the use of the family income without accounting for the individual's wage (income) variations – as in Macunovich (2011) – is not in line with Easterlin's relative income hypothesis. Aspirations per se cannot capture adequately for the latter. Macunovich was no doubt aware of this. But, she tried, despite this drawback, to conduct a more micro-oriented approach as she implies saying: "Macunovich (2002) found support for the hypothesis, however, using parental income at a more aggregated level." (Macunovich, 2011: 3).

Among the various studies dealing with Easterlin's RY hypothesis, MacDonald and Rindfuss (1981) and Macunovich (2002) appear to be best focused on the Easterlin's conceptual idea. However, their findings are contradictory: one provides evidence against the hypothesis (MacDonald and Rindfuss, 1981) whilst the other in favour of the hypothesis (Macunovich, 2002). This study contributes to the relevant literature by exploiting recent data that spans up to 2016. We also employ dynamic panel data and Granger non-causality methods that had not been considered before. We differentiate from the previous literature by incorporating the female labour force participation into the marriage model as a measure of women's emancipation (Upadhyay et al., 2014). Finally, we compare relative income to the absolute one. The comparison had also been conducted by MacDonald and Rindfuss (1981), however, without controlling for the female dimension.

2.4 Data and methods

2.4.1 Data description

The data[8] used are from the IPUMS-CPS database. IPUMS-CPS offers micro-data which we aggregate by year, state, race, and age. They span from 1976 (1981 if we account for the 5 year lags of the relative income) to 2016. The panel is unbalanced. For the RY, the numerator consists of the median wages for men aged 15-24[9] and the denominator represents the median family income with a "Head" of either sex aged 45-54. For the calculation of marriage rate, we follow Sweeney (2002) – both male and female marriage rates – and Macunovich (2011) who distinguishes between the two. Table A2.1 (Appendix) presents the descriptive statistics.

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[8] For a detailed description of the data development process, see Appendix A.

[9] In Appendix A, we note that we have taken into account unmarried young men who had worked full time, completed 52 weeks the previous year, and their calculated wage per hour was not found less than $2.50 or more than $250 in 2008 constant dollars. Both restrictions aim to exclude the possibility of educational driven results on marriage; namely, the possibility that young adults retreat from marriage due to their attendance in higher education. We test for the latter showing results for young men aged 15-24 who attained equal or more than high school and less than bachelor degree (Appendix D). The latter category has only been considered because this is the most representative educational group for the age cluster under examination (total: 43,356 individuals; less than high school: 14,726 individuals; equal or more than high school and less than bachelor degree: 25,820 individuals; equal or more than bachelor degree: 2,810 individuals). Results are in line to the ones presented in the main text.

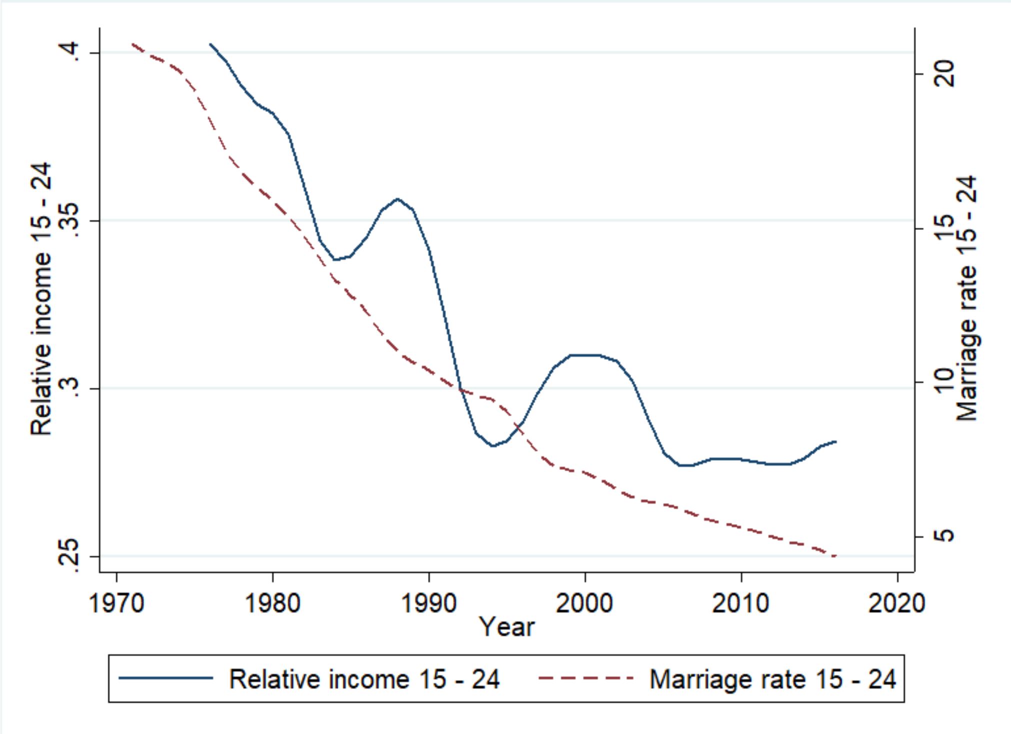

Fig. 2.4 RY of the age group 15-24 for the US. Marriage rate refers to young men (age 15-24). *Source*: Our own calculations employing data extracted from the IPUMS-CPS.

Fig. 2.4 depicts the evolution of RY and marriage for white non-Hispanics men of age 15-24 in the US. Both have been declining. For RY, we observe a decline until the middle of the 90's followed by an increase and then stagnation around 2000. After a short period of decrease, RY has been relatively steady since 2006.

2.4.2 Control variables

Other factors could also affect the decision on marriage. We consider two additional variables: female labour force participation which also serves as a measure of women's empowerment (Upadhyay et al., 2014) and the unemployment rate of young men (González-Val and Marcén, 2018). Four decade dummies have been introduced; one dummy for each decade. The use of dummies aims to capture time-varying effects which could affect marriage such as the level of inequality, cohabitation,[10] gender ratio (marriage "squeeze"), contraception, divorce rates, etc. Table A2.2 (Appendix A) provides the calculated formulas for the variables employed.

It is important to note the advantages and disadvantages of using aggregate instead of micro-data on the RY hypothesis. On the one hand, aggregate analysis arising from micro-data has been criticised when Robinson (1950) referred to the "ecological fallacy". The latter posits the incorrect assumptions made about an individual, based on the characteristics of a group to which the individual belongs. Moreover, Simpson (1951) supports the view of a reversal in the sign between analysis at the aggregate and individual-level (see also Oppenheimer, 1997).

On the other hand, tests at the micro-level have treated aspirations as being only a function of parental income by using either gross proxies for the RY measure such as "How well-off are you?" or (sparser) using the incomes of their parents. Such approaches neglect the impact of peers, neighbourhood, or other social phenomena also operates on individual aspirations. Conducting the analysis on the aggregate level, we remedy this drawback.

2.4.3 Methodology

2.4.3.1 Panel tests

We employ a panel dynamic FE estimator. The FE estimator is still consistent under unbalanced panels as long as missing observations are random (Wooldridge, 2002). Plausible estimation problems might arise from the nature of the panel data such as the correlation between and within states. Therefore, we perform tests for cross-sectional dependence (CSD), serial correlation (first order), and stationarity. The latter is also ensured by the fact that variables are bounded (Farmer, 2015) through the normalization process (see Appendix A).

[10] We can capture cohabitation indirectly by the use of dummies. IPUMS provide data on cohabiting partners from 2007 onwards.

Normalization induces a rescaling which allows one to compare the contribution between relative and absolute income. Despite the normalization process that took place, we have also employed the Pesaran (2007) unit root test for panel data too.[11]

Cross-sectional dependence has been examined with Pesaran's (2015) test. The null hypothesis of cross-sectional independence was rejected (Appendix E, Table E2.1). Lastly, we check for the presence of first order autocorrelation using the Wursten (2018) test. The presence of serial correlation would imply the loss of information hidden in the error term if we solely consider a static model.[12] That is, we would underestimate the standard errors. The inclusion of the dynamic component deals with the latter. The results shown in Appendix (Table E2.2) indicate the presence of first order autocorrelation in all cases. Hence, we conduct the dynamic model.

2.4.3.2 Model specification

We follow a dynamic approach based on the FE estimator with a lagged dependent variable. The inclusion of the lagged dependent variable may cause what Nickell (1981) has called dynamic panel bias. This could be corrected by the use of the Arellano-Bond estimator (difference GMM) which removes the fixed effects by taking first differences (Arellano and Bond, 1991; Roodman, 2009). However, the moderate value of T (41) is not small enough to avoid overfitting of the variables to be instrumented, nor large enough to argue that the dynamic panel bias is insignificant. Thus, we apply the bias-corrected dynamic panel estimator (Bruno, 2005a; 2005b) too. GMM is hence employed as a robustness check.

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[11] The null hypothesis of non-stationarity is rejected in all cases. Results are available upon request. For bounded processes unit root tests see, Carrion-i-Silvestre and Gaeda (2013), Cavaliere and Xu (2014).

[12] The static model – not presented here – finds the coefficient of the RY to be positive and statistically significant at the 1% level of significance in all cases examined. Results are available upon request.

We consider RY, female labour force participation, and young men unemployment rate as endogenous variables, since reverse causality might occur if men decide to postpone or skip marriage in order to increase their education (Nielsen et al., 2009), or to fulfill their aspirations before they get married. By increasing their education, corresponding unemployment might also be affected.[13] Thus, marriage decline could impact on RY and unemployment rate. The same might also hold for women's labour force participation. Women decide to increase their participation in the market either because men postpone (or skip) marriage, or due to their own decision to get more education and consequently higher wages. Hence, two lags have been imposed on the independent variables. The equation to be estimated is the following:

$$marriage_{i,t} = b_0 + b_1 marriage_{i,t-1} + b_2 RY_{i,t-2} + \sum_{j=3}^{4} b_j X'_{i,t-2} + \sum_{i=1}^{4} b_{5i} dummy_i + \mu_i + u_{i,t} \quad (1)$$

where $marriage_{i,t}$: marriage rate, $RY_{i,t-2}$: relative income, $X'_{i,t-2}$: control variables, $dummy_i$: decade dummy takes the value 1 for the decade of reference, otherwise takes the value 0, μ_i: fixed effect term and $u_{i,t}$: stochastic error term. To compare[14] the effect of relative to absolute income (AY henceforth) on marriage, we re-run equation (1) by replacing the RY with the AY.[15]

[13] Unemployment is mostly affected by macroeconomic factors and can be considered as exogenous to marriage; nevertheless, we have repeated the analysis without lags on unemployment. Results do not significantly deviate from the ones presented in section 2.5.1 (available upon request).

[14] Variables have been normalized so that their relative contribution on marriage is comparable.

[15] We follow Macunovich (1998a: 72) who states that: "Easterlin's hypothesis is that relative *instead of* absolute income should be used in the analysis of fertility decisions [...]. It is difficult to interpret an equation in which *both* are entered" (italics come from Macunovich).

2.4.3.3 Robustness check

We proceed to robustness check by employing three more methods. First, the Arellano-Bond estimator (difference GMM) is applied. One step difference GMM estimator has been considered instead of two[16] due to the lower standard errors obtained in the former. Moreover, we avoid the magnification of the gaps in the first difference transformation (see Roodman, 2009: 21). A control for serial correlation and heteroskedasticity has been imposed. It is also important to ensure that the number of instruments is lower than the number of states.

We also account for cross-sectional dependence (Driscoll and Kraay, 1998). The latter produces updated standard errors that correct for "spatial" forms of cross-sectional dependence. Finally, we apply the Blackwell (2005) method which corrects standard errors for heteroskedastic and contemporaneously correlated disturbances across panels.

2.4.3.4 Granger non-causality test

We apply the Granger non-Causality procedure proposed by Dumitrescu and Hurlin (2012) to test for any causal effect of relative income on marriage. Dumitrescu and Hurlin's method has the advantage to account for heterogeneous panel data models. The specification is as follows:

$$marriage_{it} = a_i + \sum_{k=1}^{K} \beta_{i,k}^{(k)} marriage_{i,t-k} + \sum_{k=1}^{K} \gamma_{i,k}^{(k)} RY_{i,t-k} + \varepsilon_{i,t} \qquad (2)$$

where k is the number of lags. The null hypothesis (no causality) is defined as:

$$H_o: \gamma_{i1} = \ldots = \gamma_{ik} = 0 \qquad \forall i = 1, \ldots, N$$

and the alternative hypothesis (causality) as:

$$H_1: \gamma_{i1} = \ldots = \gamma_{ik} = 0 \qquad \forall i = 1, \ldots, N_1$$
$$\gamma_{i1} \neq 0 \; or \ldots or \; \gamma_{ik} \neq 0 \qquad \forall i = N_1 + 1, \ldots, N$$

where$N_1 \in [0, N-1]$ is unknown.

[16] For a discussion on one and two step GMM estimator, see Hwang and Sun (2018).

This test controls for the possible difference of the unconstrained parameters from one individual (state) to another. It further takes into account the cross-sectional dependence[17] and allows the researcher to select the number of lags based on one of the selection criteria (AIC, BIC, HQIC).[18] It should be emphasized that rejecting the null hypothesis does not exclude the possibility of no causality for some states. This is an issue that should be investigated further in a future work. On the other hand, not rejecting the null hypothesis means that there is no Granger-causality for any state.

This methodology cannot be applied in unbalanced panels,[19] so we proceed with the balanced version of the dataset. The latter caused a large drop in observations (from 1724 reduced to 828). The Lopez and Weber (2017) formula has been applied to provide the relevant procedure proposed by Dumitrescu and Hurlin (2012). In the regression analysis that follows, lags have been chosen to minimize the Bayesian criterion. Finally, the variables considered for the causality test must be stationary (confirmed earlier).[20]

2.5 Estimation results

2.5.1 Regression analysis results

Table 2.1 presents the results obtained by employing the panel dynamic FE estimator. The lagged value of the dependent variable is always statistically significant at the 1% level of significance. Coefficients on both income variables are positive but statistically significant only for the RY. In all cases considered, the RY exerts a greater impact on marriage than the AY. The magnitude of the coefficient of RY and AY is higher for the "Male" category. On the covariates, both young men's unemployment and young women's labour force participation

[17] See Dumitrescu and Hurlin (2012: 3-4).

[18] See Lopez and Weber (2017: 983).

[19] See Lopez and Weber (2017: 973).

[20] See Lopez and Weber (2017: 977).

Table 2.1 Panel dynamic FE estimator.

	(1)	(2)	(3)	(4)	(5)	(6)
	Both	**Male**	**Female**	**Both**	**Male**	**Female**
Marriage_{t-1}	0.977***	0.970***	0.970***	0.981***	0.973***	0.974***
	(0.0114)	(0.00985)	(0.0125)	(0.0120)	(0.0104)	(0.0130)
RY_{t-2}	0.0286*	0.0334**	0.0283*			
	(0.0143)	(0.0158)	(0.0143)			
AY_{t-2}				0.0162	0.0185	0.0153
				(0.0131)	(0.0137)	(0.0129)
Flfp_{t-2}	-0.0773***	-0.0814***	-0.0734***	-0.0734***	-0.0764***	-0.0696***
	(0.0138)	(0.0120)	(0.0153)	(0.0136)	(0.0116)	(0.0151)
$\text{Unemployment}_{t-2}$	-0.0416***	-0.0423***	-0.0389***	-0.0436***	-0.0445***	-0.0410***
	(0.0107)	(0.0122)	(0.0110)	(0.0107)	(0.0119)	(0.0111)
Decade dummies	Included	Included	Included	Included	Included	Included
Constant	0.0598***	0.0645***	0.0642***	0.0621***	0.0681***	0.0668***
	(0.0147)	(0.0156)	(0.0156)	(0.0141)	(0.0146)	(0.0150)
R-squared	0.966	0.955	0.965	0.965	0.955	0.965
Observations	1,574	1,574	1,574	1,574	1,574	1,574

Note: dependent variable: marriage rates. *** *p-value* < 0.01, ** *p-value* < 0.05, * *p-value* < 0.1. *Flfp* stands for female labour force participation.

Table 2.2 Bias-corrected FE estimator.

	(1)	(2)	(3)	(4)	(5)	(6)
	Both	**Male**	**Female**	**Both**	**Male**	**Female**
Marriage_{t-1}	0.977***	0.970***	0.970***	0.980***	0.973***	0.974***
	(0.012)	(0.011)	(0.012)	(0.012)	(0.011)	(0.012)
RY_{t-2}	0.029***	0.033***	0.028***			
	(0.009)	(0.010)	(0.009)			
AY_{t-2}				0.016*	0.018**	0.015*
				(0.008)	(0.009)	(0.009)
Flfp_{t-2}	-0.0772***	-0.0814***	-0.073***	-0.074***	-0.076***	-0.070***
	(0.0105)	(0.011)	(0.011)	(0.010)	(0.011)	(0.011)
$\text{Unemployment}_{t-2}$	-0.042***	-0.042***	-0.039***	-0.044***	-0.045***	-0.041***
	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)
Decade dummies	Included	Included	Included	Included	Included	Included
Observations	1,574	1,574	1,574	1,574	1,574	1,574

Note: dependent variable: marriage rates. *** *p-value* < 0.01, ** *p-value* < 0.05, * *p-value* < 0.1. *Flfp* stands for female labour force participation.

are negatively related to their marriage rates. Their effect is larger in size than that of the two income variables and always statistically significant at the 1%.

Table 2.2 controls for endogeneity by employing the bias-corrected FE estimator. We observe that both incomes are positive and statistically significant (*p-value* < 0.01 for RY). As in Table 2.1, RY has a greater impact on marriage rates compared to AY. The covariates are negatively related to marriage. Lastly, female labour participation has the most profound effect.

Robustness checks are presented in Appendix (Tables B2, C2.1, C2.2, D2.1, D2.2). Results are qualitatively similar. RY coefficients are positive and statistically significant whilst significance is lost for the AY in Tables B2, D2.1, and D2.2. Male unemployment and female labour force participation are inversely related to marriage with the significance of the former being lower than the latter (Tables C2.1 and C2.2). When we compare the two incomes, we see that RY exceeds in size AY in all cases examined.

Results presented in this section (in Appendix as well) indicate that RY is an important predictor for marriage rates. Thus, the Easterlin hypothesis is corroborated. Furthermore, in all methods employed, RY is found to be more significant both in statistical and size aspects relative to the AY. Young women's empowerment leads to a decline in marriage according to findings on female labour force participation. Finally, young men's unemployment emerges as a deterrent factor on marriage.

2.5.2 Causality tests

This section employs the Dumitrescu and Hurlin (2012) Granger non-causality test as formulated by Lopez and Weber (2017). Both directions of causality have been considered (RY to marriage and marriage to RY). Table 2.3 refers to the case of the causal impact of RY on marriage. The *p-values* suggest statistical significance at the 1% level except for male marriage which lies at the 5%. Hence, we can reject the null hypothesis of no-causality and infer causality for at least one state in all cases presented.

Table 2.4 tests for the causal impact of marriage on RY. The *p-values* change and become higher for "Both" and "Female", but still low in the case of "Male". Thus, results indicate that there is a causal effect of marriage on relative income only for young men ("Male"). This might be explained by the fact that young men who retreat from marriage have lower relative income than their peers who get married (Fig. 2.5). In sum, results show that in five out of six cases considered, the causal effect is one-directional and runs from RY to marriage.

Table 2.3 Dumitrescu and Hurlin (2012) Granger non-causality test.

	Both	Male	Female
W-bar (Wald statistic)	31.3014	31.8950	26.9472
Z-bar (Standardized statistic)	32.9531	33.7926	26.7954
p-value[a]	0.0080	0.0330	0.0010
Critical value (95%)	105.5197	87.0666	58.3001
Z-bar tilde[b]	9.7544	10.0425	7.6415
p-value[a]	0.0080	0.0330	0.0010
Critical value (95%)	34.6543	28.3225	18.4518
Optimal number of lags[c]	8	8	8

H_0:RY does not Granger-cause marriage.
H_1:RY does Granger-cause marriage for at least one state.
[a]*p-value* computed using 1000 bootstrap replications.
[b]Preferred for large N but relatively small T datasets.
[c]Optimal number of lags selected according to Bayesian (BIC) criterion.

Table 2.4 Dumitrescu and Hurlin (2012) Granger non-causality test.

	Both	Male	Female
W-bar (Wald statistic)	30.7712	38.5508	24.0406
Z-bar (Standardized statistic)	32.2034	43.2053	22.6849
p-value[a]	0.1520	0.0270	0.5630
Critical value (95%)	38.0339	39.1517	40.7129
Z-bar tilde[b]	9.4972	13.2723	6.2311
p-value[a]	0.1520	0.0270	0.5630
Critical value (95%)	11.4978	11.8813	12.4171
Optimal number of lags[c]	8	8	8

H_0:Marriage does not Granger-cause RY.
H_1:Marriage does Granger-cause RY for at least one state.
[a]p-value computed using 1000 bootstrap replications.
[b]Preferred for large N but relatively small T datasets.
[c]Optimal number of lags selected according to Bayesian (BIC) criterion.

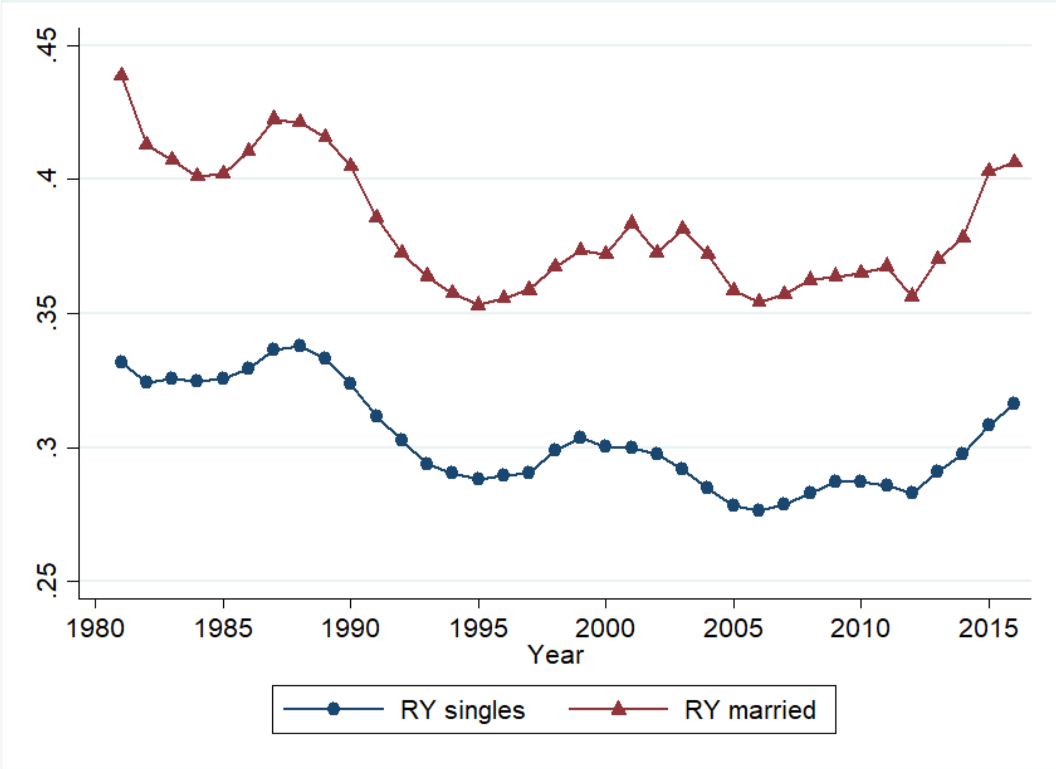

Fig. 2.5 Relative income for single and married young men (age 15-24). *Source*: Our own calculations employing data extracted from the IPUMS-CPS.

2.6 Conclusions

Previous studies have indicated the importance of the institution of marriage for individual and society's well-being (Popenoe, 2009; Lerman, 2011; Halla and Scharler, 2012). This paper investigates its decline.

We reveal that an increase in the relative income of young men increases their odds to wed. Hence, there is some space for the policy maker to intervene. RY could be affected either by the numerator (income or wage) or the denominator (childhood income). The former could be accommodated through a stabilising macroeconomic policy. For the latter, one could claim the same because the current income of parents is the future childhood income of young adults. However, the impact of a wage increase would affect (childhood) income inequality which in turn would be depicted in future RY (in young adulthood). Thus, one has to consider the impact of income inequality in a long-run perspective.[21] Richins (2017) found that middle school children are especially engaged in social comparisons, and this is why childhood inequality becomes important. Therefore, the higher the income inequality, the worst the outcome of the comparisons for the less advantaged children, and the higher the materialism the latter acquire and transfer to their early adulthood. Hence, one could primarily infer that RY indicates the necessity for a policy implementation to reduce income inequality. This might reduce materialism levels arising from this source and lower the consumption threshold of young men.

To conclude, the results we present indicate that policy makers should also regard aspirations along with income. We infer that neither the current economic position of young adults per se, nor their aspirations per se, but their ratio is one of the key variables on marriage decision.

2.7 References

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[21] See also Genicot and Ray (2017).

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2.8 Appendix

A. Data derivation/construction

Data begin in 1962. We have, however, constrained time dimension due to the lack of the relevant variables needed for the development of RY. Consequently, the analysis covers the period between 1976 (1981 if we account for the 5 year lags of the relative income) and 2016, taking place in a year-state-race-age level basis. The States included are the following: Alabama, Arizona, Arkansas, California, Colorado, Connecticut, Delaware, Florida, Georgia, Idaho, Illinois, Indiana, Iowa, Kansas, Kentucky, Louisiana, Maine, Maryland, Massachusetts, Michigan, Minnesota, Mississippi, Missouri, Montana, Nebraska, Nevada, New Hampshire, New Jersey, New Mexico, New York, North Carolina, North Dakota, Ohio, Oklahoma, Oregon, Pennsylvania, Rhode Island, South Carolina, South Dakota, Tennessee, Texas, Utah, Vermont, Virginia, Washington, West Virginia, Wisconsin, Wyoming. The analysis has been constrained to white non-Hispanics due to their large representativeness in the sample of the IPUMS-CPS database. For instance, the main file of IPUMS-CPS downloaded contains 8,677,348 observations of which 7,280,335 refer to whites and 6,674,636 (out of 8,677,348) to white non-Hispanics. The sample is also constrained to civilians as in Macunovich (2012). All variables have been estimated by year, state, race, and age.

The numerator of the developed RY consists of the wages for men of age between 15 and 24 years old, who are singles (never married), worked full time, and completed 52 weeks the previous year (full time-year workers). Wages refer to total pre-tax wage and salary income for the previous calendar year. It is crucial to include only those who have completed 52 weeks, for annual wages earned by individuals would present considerable heterogeneity and hence would not be comparable.

We also restricted the sample to those who are native born (we use the variable "bpl" – birth place – which unfortunately starts since 1994) in order to increase the odds to capture their aspirations in childhood. Next, we replaced the top-coded wages with the swap values for the years 1976-2010 and dropped for the rest of the period (see https://cps.ipums.org/cps /topcodes_tables.shtml). Moreover, we dropped individuals from the analysis if their calculated wage per hour was less than $2.50 or more than $250 in 2008 constant dollars (following Blau and Kahn, 2007). The wage was estimated as the median of all individuals by year, state, race, and age. We considered the median calculation instead of the mean because the wage distribution was mostly skewed toward the right side. In the denominator, we calculated the family income which has been lagged by five years in order to proxy the aspirations of young males. The choice of five year lags (instead of three or ten for instance) relates to Easterlin's initial work (see Easterlin, 1966; Macunovich, 1998a) and data limitations. By the former we stay closer to Easterlin's pure approach and by the latter we save observations which are already limited. For instance, if we pick a lag of three years, the max age of the individuals considered in the analysis, that is, 24 years old, would relate to aspirations when they were 21 years old (high enough). On the other hand, a lag choice of 10 years would catch aspirations plausibly better than 5, but with the cost of fewer observations.

The use of family income (total family income) instead of male's is justified by the increased pace of women's contribution to family income and therefore on child's aspirations (Macunovich, 2011). By the selection of the specific age group of young adults (15-24) and the assumed mean age at birth (30 years old), we conclude that their parents' age will range between 45 and 54 (15-24, +30) where the parental family income has been calculated and lagged. We adopt the age of thirty since we aim to capture aspirations for all possible birth orders of the young adults. For male's mean age at birth in the US see Khandwala et al. (2017) and for female's see Mathews and Hamilton (2002, 2016).

In contrast to the numerator of RY, the denominator has not been restricted to natives born. Even if the individuals are immigrants, or of a lower socio-economic status, they contribute equally to the median family income of each state which is assumed to form the aspirations of children and teens; psychologists have since a long time indicated the tendency of individuals on social comparisons (Festinger, 1954), which could be either upward or downward (Buunk & Gibbons, 2007). Accordingly, we consider all categories in the denominator with no exclusions, in order to account for both directions. We kept in the analysis only those who reported having at least one child and be the householder (coded as "Head"). Again, as in the numerator, we estimated the median instead of the mean for the reason already mentioned. Finally, variables have been smoothed[22] and normalized (sum to unity) to address scale disparity.

Table A2.1 Descriptive statistics.

Variable	Obs.	Mean	Std. Dev.	Min	Max
Marriage rate (b)	1,724	13.452	7.374	0.754	38.462
Marriage rate (f)	1,724	15.728	8.522	0.714	43.421
Marriage rate (m)	1,724	8.944	5.447	0.465	28.049
Relative Income	1,724	0.308	0.072	0.108	0.915
Absolute Income	1,724	25010.48	4827.085	10000.08	47297.54
Female labour force participation	1,724	56.114	8.829	28.283	82.353
Unemployment rate (m)	1,724	13.503	5.812	1.754	38.889

Note: *f*: female, *m*: male, *b*: both males and females.

Table A2.2 Calculation formulas for the employed variables.

Variable	Calculation[a]
Marriage rate	$\sum_{i=1}^{48}\sum_{t=1981}^{2016}\frac{\text{Number of married}_{i,t}}{\text{Total population}_{i,t}} * 100$
Relative Income	$\sum_{i=1}^{48}\sum_{t=1981}^{2016}\frac{\text{Male wage}_{i,t}}{\text{Family income}_{i,t-5}}$
Absolute Income	$\sum_{i=1}^{48}\sum_{t=1981}^{2016}\text{Male wage}_{i,t}$
Labour force participation of young women	$\sum_{i=1}^{48}\sum_{t=1981}^{2016}\frac{\text{Female labour force}_{i,t}}{\text{Total populat. of women}_{i,t}} * 100$
Unemployment rate of young men	$\sum_{i=1}^{48}\sum_{t=1981}^{2016}\frac{\text{Number of unempl. men}_{i,t}}{\text{Lab. force partic. of men}_{i,t}} * 100$

Note: [a]*i* refers to state and *t* to year. In all cases, we adjust for white non-Hispanics of the age group 15-24. Data extracted from the IPUMS-CPS.

[22] Smoother specification has been based on running medians.

B. GMM estimator results

Table B2 GMM first difference estimator.

	(1) Both	(2) Male	(3) Female	(4) Both	(5) Male	(6) Female
Marriage $_{t-1}$	1.519***	1.100***	1.238***	1.718***	1.011***	1.408***
	(0.341)	(0.290)	(0.351)	(0.243)	(0.276)	(0.242)
RY $_{t-2}$	0.929**	0.817*	0.846*			
	(0.461)	(0.412)	(0.459)			
AY $_{t-2}$				0.625	0.317	0.552
				(0.390)	(0.614)	(0.388)
Flfp $_{t-2}$	-1.221***	-0.933***	-1.076***	-0.885***	-0.604***	-0.777***
	(0.219)	(0.195)	(0.234)	(0.205)	(0.158)	(0.202)
Unemployment $_{t-2}$	-1.076***	-1.223***	-1.171***	-0.905***	-1.278***	-1.028***
	(0.311)	(0.325)	(0.304)	(0.239)	(0.309)	(0.245)
Decade dummies	Included	Included	Included	Included	Included	Included
F-stat. (Prob>F)	0.000	0.000	0.000	0.000	0.000	0.000
Number of instruments	41	41	41	41	41	41
Observations	1,508	1,508	1,508	1,508	1,508	1,508

Note: Dependent variable: marriage rates. Instrumental variable: one year lag of the independent. *** *p-value*< 0.01, ** *p-value*< 0.05, * *p-value*< 0.1. *Flfp* stands for female labour force participation.

C. Control for cross-sectional dependence (CSD)

Table C2.1 Control for CSD (D-K s.e.).

	(1) Both	(2) Male	(3) Female	(4) Both	(5) Male	(6) Female
Marriage $_{t-1}$	0.977***	0.970***	0.970***	0.981***	0.973***	0.974***
	(0.0210)	(0.0220)	(0.0203)	(0.0227)	(0.0234)	(0.0217)
RY $_{t-2}$	0.0286**	0.0334***	0.0283*			
	(0.0137)	(0.0112)	(0.0155)			
AY $_{t-2}$				0.0162**	0.0185**	0.0153*
				(0.00786)	(0.00876)	(0.00885)
Flfp $_{t-2}$	-0.0773**	-0.0814***	-0.0734**	-0.0734**	-0.0764***	-0.0696**
	(0.0322)	(0.0263)	(0.0341)	(0.0310)	(0.0254)	(0.0330)
Unemployment $_{t-2}$	-0.0416	-0.0423*	-0.0389	-0.0436*	-0.0445*	-0.0410
	(0.0248)	(0.0230)	(0.0255)	(0.0255)	(0.0233)	(0.0263)
Decade dummies	Included	Included	Included	Included	Included	Included
Constant	0.0598***	0.0645***	0.0642***	0.0621***	0.0681***	0.0668***
	(0.0142)	(0.0161)	(0.0152)	(0.0141)	(0.0161)	(0.0148)
R-squared	0.955	0.963	0.965	0.955	0.962	0.965
Observations	1,574	1,574	1,574	1,574	1,574	1,574

Note: Dependent variable: marriage rates. Discroll-Kraay s.e. in parentheses. *** *p-value*< 0.01, ** *p-value*< 0.05, * *p-value*< 0.1. *Flfp* stands for female labour force participation.

Table C2.2 Control for CSD (*xtpcse*).

	(1) Both	(2) Male	(3) Female	(4) Both	(5) Male	(6) Female
Marriage$_{t-1}$	0.956***	0.942***	0.957***	0.974***	0.959***	0.973***
	(0.0314)	(0.0263)	(0.0316)	(0.0319)	(0.0263)	(0.0321)
RY$_{t-2}$	0.0508**	0.0534**	0.0499**			
	(0.0236)	(0.0232)	(0.0235)			
AY$_{t-2}$				0.0432**	0.0432**	0.0389*
				(0.0198)	(0.0187)	(0.0199)
Flfp$_{t-2}$	-0.0802***	-0.0772***	-0.0828***	-0.0824***	-0.0783***	-0.0844***
	(0.0267)	(0.0242)	(0.0280)	(0.0271)	(0.0245)	(0.0284)
Unemployment$_{t-2}$	-0.0669***	-0.0673***	-0.0669***	-0.0664***	-0.0651***	-0.0663***
	(0.0258)	(0.0238)	(0.0256)	(0.0255)	(0.0232)	(0.0255)
Decade dummies	Included	Included	Included	Included	Included	Included
Constant	0.0717*	0.0771**	0.0755*	0.0617	0.0678*	0.0675
	(0.0395)	(0.0335)	(0.0410)	(0.0424)	(0.0360)	(0.0437)
R-squared	0.937	0.928	0.938	0.937	0.928	0.938
Observations	1,574	1,574	1,574	1,574	1,574	1,574

Note: Dependent variable: marriage rates. *** *p-value*< 0.01, ** *p-value*< 0.05, * *p-value*< 0.1. *Flfp* stands for female labour force participation.

D. Results for the middle educated young men (equal or more than high-school and less than bachelor degree)

Table D2.1 Panel dynamic FE estimator.

	(1) Both	(2) Male	(3) Female	(4) Both	(5) Male	(6) Female
Marriage$_{t-1}$	0.891***	0.907***	0.884***	0.899***	0.910***	0.892***
	(0.0142)	(0.0176)	(0.0147)	(0.0139)	(0.0170)	(0.0146)
RY$_{t-2}$	0.0401**	0.0325*	0.0474**			
	(0.0189)	(0.0171)	(0.0214)			
AY$_{t-2}$				0.00361	0.00146	0.0115
				(0.0108)	(0.00885)	(0.0121)
Flfp$_{t-2}$	-0.00861	-0.0176	-0.00552	-0.00740	-0.0164	-0.00396
	(0.0116)	(0.0112)	(0.0132)	(0.0114)	(0.0110)	(0.0130)
Unemployment$_{t-2}$	0.00809	0.0161**	0.00664	0.00565	0.0139*	0.00409
	(0.00725)	(0.00731)	(0.00753)	(0.00750)	(0.00730)	(0.00796)
Constant	0.00557	-0.00343	0.0102	0.0171	0.00906	0.0199
	(0.0149)	(0.0179)	(0.0153)	(0.0161)	(0.0173)	(0.0166)
Decade dummies	Included	Included	Included	Included	Included	Included
Observations	1,059	1,059	1,059	1,059	1,059	1,059

Note: Dependent variable: marriage rates. *** *p-value*< 0.01, ** *p-value*< 0.05, * *p-value*< 0.1. *Flfp* stands for female labour force participation.

Table D2.2 Bias-corrected FE estimator.

	(1) Both	(2) Male	(3) Female	(4) Both	(5) Male	(6) Female
$Marriage_{t-1}$	0.891***	0.907***	0.884***	0.899***	0.910***	0.892***
	(0.015)	(0.014)	(0.015)	(0.015)	(0.014)	(0.015)
RY_{t-2}	0.040***	0.033***	0.048***			
	(0.013)	(0.012)	(0.014)			
AY_{t-2}				0.004	0.002	0.012
				(0.010)	(0.009)	(0.011)
$Flfp_{t-2}$	-0.009	-0.018*	-0.006	-0.007	-0.017*	-0.004
	(0.010)	(0.010)	(0.011)	(0.014)	(0.010)	(0.011)
$Unemployment_{t-2}$	0.008	0.016**	0.007	0.006	0.014**	0.004
	(0.007)	(0.010)	(0.007)	(0.007)	(0.006)	(0.007)
Decade dummies	Included	Included	Included	Included	Included	Included
Observations	1,059	1,059	1,059	1,059	1,059	1,059

Note: Dependent variable: marriage rates. *** *p-value*< 0.01, ** *p-value*< 0.05, * *p-value*< 0.1. *Flfp* stands for female labour force participation.

E. Panel tests

Table E2.1 Pesaran (2015) cross-sectional dependence test.

Variable	CD-test	*p*-value	Average joint T	Mean ρ	Mean abs (ρ)
Marriage(b)	178.121	0.000	34.10	0.91	0.91
Marriage(m)	166.213	0.000	34.10	0.85	0.85
Marriage(f)	178.597	0.000	34.10	0.91	0.91
Relative Income	74.157	0.000	34.10	0.38	0.43
Absolute Income	26.631	0.000	34.10	0.13	0.31
Flfp	144.385	0.000	34.10	0.74	0.74
Unemployment	64.245	0.000	34.10	0.33	0.42

H_0:cross-section independence.

Note: *f*: female, *m*: male, *b*: both males and females.

Table E2.2 Wursten (2018) first order serial correlation test.

Case	Variable	HR-stat[a]	*p*-value
Marriage(b)	Residual	9.06	0.000
Marriage(m)	Residual	6.42	0.000
Marriage(f)	Residual	9.06	0.000

H_0:no first-order serial correlation.

[a]Heteroskedasticity-robust Born and Breitung (2016) HR-test on residuals.

Note: *f*: female, *m*: male, *b*: both males and females.

CHAPTER 3

The Increasing Rate of Children born Out-of-Wedlock in the US

Abstract

This paper investigates the evolution of non-marital fertility in the United States during the last decades. We examine the impact of the material aspirations formed in childhood compared to the current economic situation of young unmarried adults on births out-of-wedlock. This is the Easterlin relative income hypothesis, which has been overlooked by the relevant literature. Panel data has been employed and covers the period from 1981 to 2016. Evidence suggests a negative and statistically significant effect of relative income on the fraction of children born out-of-wedlock. Most important, results are robust to the inclusion of marriage which might imply a socio-economic mobility perspective. We proceed to a comparison to the effect of the absolute income. Relative income emerges as a stronger predictor than the latter in most of the cases examined.

Keywords: non-marital fertility, relative income, Easterlin hypothesis.

JEL Classification Codes: J12, J13, J19.

3.1 Introduction

The increased rate of children born out-of-wedlock has been one of the main demographic features during the past decades in the developed world. Van de Kaa (1987) introduced the term "Second Demographic Transition" in order to describe this phenomenon, along with the attenuation of marriage and the trend to higher cohabitation rates. United States (US) has experienced an unprecedented rise of non-marital fertility rates. In 1960, birth rate per 1000 unmarried women aged 15-44 was 21.6 and increased to 43.4 in 2015 – reaching its max (51.8) in 2007 (Fig. 3.1). In 2015, birth rate per 1000 unmarried women was 67.4 for Hispanics, 59.6 for blacks, 40 for whites, and 31.6 for white non-Hispanics. Except for race differences, non-marital fertility in the US varies by age and educational level too.

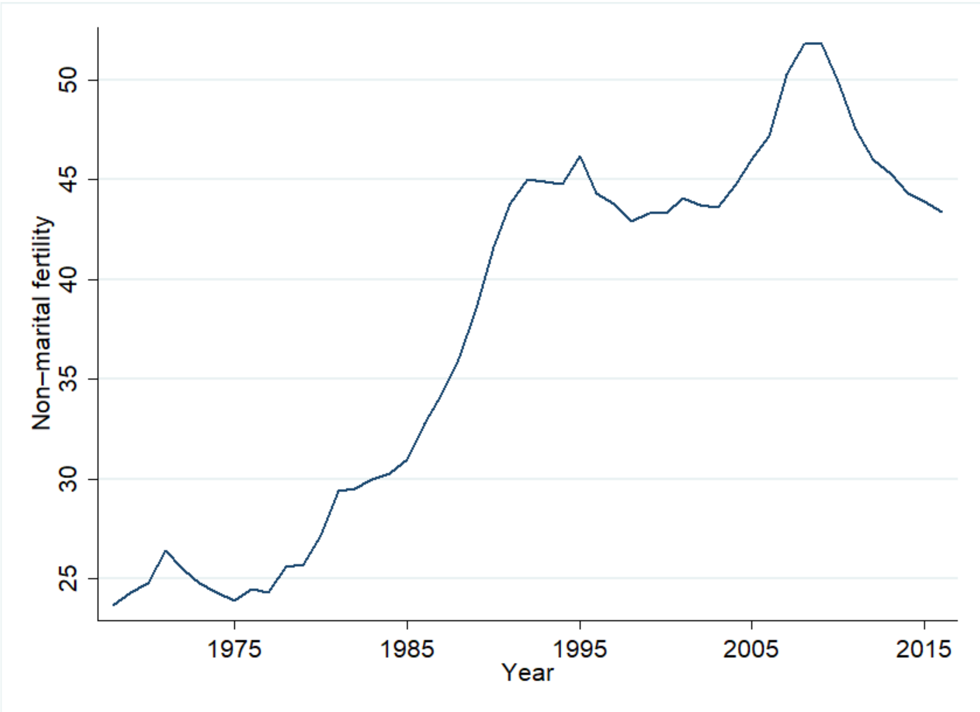

Fig. 3.1 Birth rates to unmarried women of age 15-44 per 1000 unmarried women in the US. *Source*: Constructed by the Author employing data from Hamilton et al. (2015: 42).

Among teens between 15 and 19 years old, the highest rate in 2015 was recorded for Hispanics (31.6) and the lowest one for white non-Hispanics (13.9). There is also some evidence that non-marital fertility coincides with negative educational gradient (Rindfuss et al., 1996; Sweeney, 2002). In particular, Kearney and Levine (2017) found that 62% among women with a high school or less educational attainment have a child outside marriage. Despite the decrease observed the last decade, the fraction of children born out-of-wedlock remains at high levels (Kearney and Levine, 2012), which raises some concerns over well-being repercussions (Bellido et al., 2016; Liang et al., 2019).

This paper aims to show that one of the drivers on non-marital fertility is the relative deprivation of young men. To our knowledge, Easterlin's hypothesis has never been tested before for premarital births. We test for the latter focusing on white non-Hispanics for the period 1981-2016 in the US. Results support the hypothesis.

3.2 Literature review

There are many theoretical and empirical approaches to the non-marital fertility phenomenon in the US. Theoretical papers emphasize unemployment (Wilson, 1987) and changes in welfare incentives (Murray, 1984). Akerlof et al. (1996) argue that such interpretations can only explain a small fraction of the change. They claim that the increase of non-marital fertility in past decades comes from the decline of "shotgun marriage" caused by the availability/cost of abortion, spread of contraception[1], and lower social isolation. Another reason might be welfare provision to single mothers. However, the welfare effect hypothesis has not been adequately supported by the literature. Some scholars find a negative effect (Licther et al., 1997; Garfinkel, 2003) whilst others indicate

[1] The spread of contraception and availability/cost of abortion could lead to two opposite effects: first, it would reduce non-marital fertility due to their usage, and second, it would increase odds to non-marital fertility because of a higher premarital sexual activity (moral hazard); if the latter exceeds the former, births out-of-wedlock will possibly increase.

a positive one (Moffitt, 1992; 1994). Overall, welfare payments have a weak impact on non-marital fertility (Moffitt, 2001).

On job availability, Olsen and Farkas (1990) claim that underclass youth employment opportunities increase the odds to family formation, and Schneider (2015) finds a significant negative relationship between state-level unemployment and non-marital fertility for women, independent of their socio-economic status.

Becker's traditional economic model of marriage (Becker, 1991) suggests that the out-of-wedlock childbearing becomes prevalent when women possess the sufficient income to establish their own family. The latter may be reinforced by the reduced gains from marriage due to the low income of young men. In the same vein, Willis (1999) shows that women prefer out-of-wedlock childbearing if their income is high enough. Furthermore, Boschini et al. (2011) find a positive correlation between education and childlessness among women, in contrast to that observed in the case of men (Lappegård, 2011). Aassve (2003) argues that the increased female wage has only a small effect owing to its contradictive results on fertility and marriage. Using the predicted wage of women, he demonstrates that when wage goes up, only a few women either get married or born children out-of-wedlock found that a higher income of men reduces the time to marriage implying a reduction on non-marital births. They also show that the effect of women's income is to delay marriage.

Becker and Willis, as presented above, imply that relative income (male to female wage) stands at the center of interest rather than the absolute income per se. Similarly, Bertrand et al. (2015) present evidence that partners prefer matches in which earnings of men exceed that of women. Except for the argument on gender relative earnings, research on non-marital fertility has also mentioned the impact of income with respect to a reference group (Watson and McLanahan, 2011). Finally, Ellwood and Jencks (2004) concentrate on some non-economic explanations including gender role conflict, limited confidence and personal efficacy, altered attitudes, social norms, as well as technological and legal change.

However, non-marital fertility literature has hitherto overlooked the Easterlin relative income hypothesis. To the best of our knowledge, the latter has hardly ever tested within a non-marital fertility framework. We draw data from the IPUMS-CPS database for the period 1981-2016 for the US. Empirical results show that relative income is negative and statistically significant to the fraction of children born out-of-wedlock after controlling for marriage and other covariates. We contribute to the non-marital fertility literature indicating the Easterlin relative income as one of the determinants on the evolution of children born out-of-wedlock during the last 36 years in the US.

3.3The relative income hypothesis

Easterlin claims that young people's behaviour is determined by relative instead of absolute income. An important issue is to clarify what Easterlin means by saying relative income. In a conversation with Diane J. Macunovich, Easterlin stated: "By relative income, I mean one's assessment of earnings prospects in relation to an internalized norm of one's desired living level – what one might think of as a socially defined subsistence level" (Macunovich, 1997: 121). Therefore, according to Easterlin, young adults will decide on various aspects of their life (fertility among others) by comparing their current earnings prospects with their desired level of living. If they feel that their earnings are high enough compared to their desired standard of living, they will marry, childbearing etc. On the opposite, if their earnings are low (or their standard of living high, or both), they will feel relative deprived and hence avoid or postpone marriage and childbearing. There is, therefore, a minimum consumption threshold.

Relative income hypothesis is highly dependent on the way one proxies the minimum consumption threshold of young adults – the denominator of relative income (Macunovich, 1997). Easterlin (1966) employed the family income of a specific age group as a proxy to the subsistence level. He considers that material aspirations (henceforth aspirations) are formed during childhood according to family income levels and play a key role in early adulthood decision process. The relation is assumed to be positive.[2] Thus, higher levels of family

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[2] This is the relation adopted in the fertility literature. However, Easterlin, in his effort to solve the happiness paradox and based on surveys for the US, claims that aspirations are fairly

income will produce higher levels of aspirations and given the current absolute income (assumed constant) will depress relative income. This point needs caution. By taking family income as a proxy for the subsistence level of young adults during childhood, would be false to infer that most young adults expect to earn in their early adulthood what their parents had earned some years earlier. This misleading view along with the import of categorical variables of surveys in the denominator has brought ambiguity on Easterlin's hypothesis (Maunovich, 1998). The proxy implies that the minimum consumption threshold for young adults will be a *function* of family income. Hence, family income could show us the degree to which aspirations interacting with current absolute income affect the decision process for young adults.

3.4 Data and methodology

Data retrieved from the IPUMS-CPS and spans the period 1981-2016 for the US.[3] In brief, relative income represents a ratio. The numerator includes the incomes of young single adults aged 15-24 years old. The denominator corresponds to the family income of the "Heads" aged 45-54 years old lagged by five years. The latter has been used as a proxy for childhood material aspirations (Easterlin, 1966).

We employ a dynamic panel data fixed effect (FE) estimator. FE estimator could deal with some of the endogeneity problems as it controls for all the time-invariant variables. We try to capture time-varying effects by the inclusion of decade dummies. Robust standard errors aim to control for autocorrelation and heteroscedasticity. Dynamic panel data FE with Discroll-Kraay standard errors (Discroll and Kraay, 1998) has been further performed to control for the presence of cross-sectional dependence.

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similar among different income groups of young adults due to peer influence in childhood. But the peer influence attenuates over the life cycle and "social comparison influences are increasingly confined to a reference group comprised of those of one's own socio-economic status", Easterlin (2001: 480).

[3] For a detailed description of the data development process, see Appendix B.

In order to deal with endogeneity induced by the dynamic component (Nickell, 1981), we employ the Arellano-Bond difference GMM estimator (Arellano and Bond, 1991; Roodman, 2009). In the absence of any other accurate instrumental variable for relative income, its lagged value serves as an instrument. Due to the moderate value of T (35)[4] we also consider the bias-corrected dynamic panel estimator (Bruno, 2005a; 2005b).

Finally, panel tests for stationarity (Pesaran, 2007), first order serial correlation (Wursten, 2018), and for cross-sectional dependence (Pesaran, 2015) have been conducted. In Appendix A, we see that all variables are stationary (Table A3.1); the null hypothesis of no first order serial correlation cannot be rejected (Table A3.3), and we also find that we are under the effect of cross-sectional dependence (Table A3.2).

The specification is as follows.[5]

First, on the relative income (RY):

$$nmf_{i,t} = b_0 + b_1 nmf_{i,t-1} + b_2 RY_{i,t-2} + \sum_{j=3}^{7} b_j X'_{i,j,t-k} + \sum_{i=1}^{4} b_{8i} dummy_i + \mu_i + u_{i,t} \quad (1)$$

Second, on the absolute income (AY):

$$nmf_{i,t} = b_0 + b_1 nmf_{i,t-1} + b_2 AY_{i,t-2} + \sum_{j=3}^{7} b_j X'_{i,j,t-k} + \sum_{i=1}^{4} b_{8i} dummy_i + \mu_i + u_{i,t} \quad (2)$$

where $nmf_{i,t}$ denotes the fraction of children born out-of-wedlock, $RY_{i,t-2}$ is the relative income lagged by two years, $AY_{i,t-2}$ is the absolute income lagged by two years, $X'_{i,j,t-k}$ is a vector of control variables (see Appendix C), μ_i is the fixed-effect term, and $u_{i,t}$ is the stochastic error term.

[4] The number of T is neither small enough to avoid overfitting the variables to be instrumented, nor large enough to argue that the dynamic panel bias is insignificant.

[5] See footnote 15, p. 28, for an explication.

3.5 Results

In Fig. 3.2 we present the obtained relationship between non-marital fertility and relative income. The relationship is negative and well fitted.

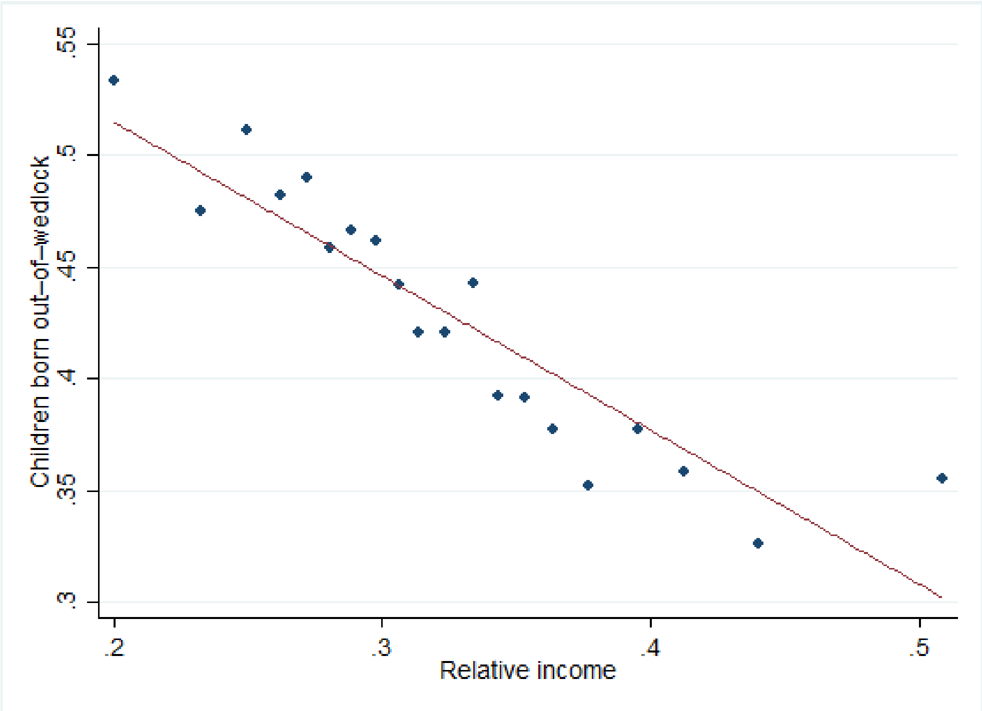

Fig. 3.2 The relationship between relative income and non-marital fertility for young adults aged 15-24. *Source*: Author's calculations employing data extracted from IPUMS-CPS.

Tables 3.1 and 3.2 present the results obtained from the regression analysis. We observe that both Tables provide evidence on a negative and – in most cases – statistically significant effect of relative income on the fraction of children born out-of-wedlock. Relative income has a greater impact on *nmf* than the absolute one, apart from one case in the GMM estimator.

In particular, in Table 3.1, relative income is negative and statistically significant at the 1% level of significance for the dynamic FE estimator (column 1). It becomes insignificant when other variables are included (column 2). We notice the same reaction for the absolute income (columns 3 and 4). However, when we control for cross-sectional dependence, using the D-K s.e. formula, we observe that the relative income continues to be statistical significant even after the inclusion of the control variables, albeit at the 10% level of significance (column 6). On the other hand, absolute income remains insignificant (column 8).

In Table 3.2, the picture is slightly different. Both incomes are again significant at the 1% level of significance for the bias-corrected FE estimator. In contrast to Table 3.1, they keep their significance at the same level when covariates are incorporated into the model. The latter does not hold in case of the GMM estimator (columns 5 - 8). Thus, in columns 6 and 8, significance is lost for both incomes and this is the only case where the absolute income reveals a stronger impact compared to the relative one. Yet, we have to mention that the insignificance of the dynamic dependent variable (nmf_{t-1}) in columns 6 and 8 make us even more hesitant on the results obtained applying the GMM estimator.

In both Tables, independent of the statistical significance, relative income is always negative which confirms our main hypothesis. The latter means that an increase of young men's relative income induces a reduction on the children born out-of-wedlock. Notice that the aforementioned result is – in two cases – robust to the inclusion of marriage. That is, relative income exerts a "marriage independent" effect on the observed non-marital fertility evolution; not only through marriage rates (Macunovich, 2011). One could easily attribute the investigated relationship (*nmf* – relartive income) to the decline of marriage rates: marriage retreats and hence non-marital fertility increases. Thus, the relationship under examination could be thought as a spurious one. By the inclusion of marriage rates into the regression, the latter collapses. We comment further in the conclusion.

Table 3.1 Dynamic FE and dynamic with Discroll-Kraay s.e. FE estimators.

Dependent variable:	Dynamic FE				Dynamic FE (D-K s.e.)			
nmf	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
nmf_{t-1}	0.920***	0.861***	0.937***	0.864***	0.920***	0.861***	0.937***	0.864***
	(0.00873)	(0.0162)	(0.00550)	(0.0156)	(0.0414)	(0.0274)	(0.0390)	(0.0274)
Relative $income_{t-2}$	-0.0588***	-0.0183			-0.0588**	-0.0183*		
	(0.0199)	(0.0199)			(0.0219)	(0.00961)		
Absolute $income_{t-2}$			-0.050***	-0.0015			-0.050**	-0.0015
			(0.0167)	(0.0163)			(0.0215)	(0.00854)
$Marriage_{t-2}$		-0.236***		-0.239***		-0.236***		-0.239***
		(0.0314)		(0.0318)		(0.0764)		(0.0775)
$Flfp_{t-2}$		-0.0447**		-0.0475**		-0.0447		-0.0475
		(0.0195)		(0.0199)		(0.0291)		(0.0284)
Female $wage_{t-2}$		-0.00355		-0.00605		-0.00355		-0.00605
		(0.0180)		(0.0178)		(0.0147)		(0.0128)
$Premium_{t-2}$		0.0398*		0.0387*		0.0398**		0.0387*
		(0.0220)		(0.0217)		(0.0192)		(0.0194)
Male $unemployment_{t-2}$		0.00752		0.00739		0.00752		0.00739
		(0.0163)		(0.0164)		(0.0120)		(0.0120)
Decade dummies	No	Yes	No	Yes	No	Yes	No	Yes
Constant	0.0703***	0.193***	0.0573***	0.187***	0.0703***	0.193***	0.0573***	0.187***
	(0.0125)	(0.0298)	(0.00855)	(0.0288)	(0.0217)	(0.0533)	(0.0193)	(0.0520)
R-squared	0.933	0.946	0.932	0.945	0.933	0.946	0.932	0.945
Observations	1,521	1,521	1,521	1,521	1,521	1,521	1,521	1,521

Note: Robust standard errors in parentheses. *** *p–value* < 0.01, ** *p–value* < 0.05, * *p–value* < 0.1. *Flfp* stands for female labour force participation. *nmf* stands for children born out-of-wedlock (non-marital fertility). Columns 1 - 4 refer to dynamic FE and columns 5 - 9 refer to dynamic FE with Discroll-Kraay standard errors.

Table 3.2 Bias-corrected dynamic FE and GMM difference estimator.

Dependent variable:	**Dynamic FE (bias corrected)**				**GMM difference**[a]			
nmf	**(1)**	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**	**(7)**	**(8)**
nmf_{t-1}	0.913***	0.824***	0.948***	0.825***	0.383**	0.387	0.553***	0.253
	(0.008)	(0.012)	(0.070)	(0.012)	(0.178)	(0.277)	(0.157)	(0.246)
Relative $income_{t-2}$	-0.123***	-0.054***			-1.019***	-0.245		
	(0.011)	(0.011)			(0.326)	(0.551)		
Absolute $income_{t-2}$			-0.100***	-0.039***			-0.572**	-0.434
			(0.011)	(0.011)			(0.220)	(0.561)
$Marriage_{t-2}$		-0.274***		-0.279***		-0.815**		-0.871***
		(0.018)		(0.018)		(0.360)		(0.309)
$Flfp_{t-2}$		-0.035**		-0.346**		0.272		0.0881
		(0.015)		(0.015)		(0.384)		(0.345)
Female $wage_{t-2}$		-0.004		-0.003		-0.748		-0.530
		(0.011)		(0.011)		(0.467)		(0.426)
$Premium_{t-2}$		0.029*		0.027*		1.015		0.895
		(0.015)		(0.015)		(1.024)		(1.114)
Male $unemployment_{t-2}$		-0.015		-0.011		0.0608		0.0621
		(0.010)		(0.010)		(0.223)		(0.226)
Decade dummies	No	Yes	No	Yes	No	Yes	No	Yes
F-stat. (Prob>F)					0.000	0.000	0.000	0.000
Number of instruments					32	35	32	35
Observations	1,591	1,591	1,591	1,591	1,453	1,453	1,453	1,453

Note: *** *p–value* < 0.01, ** *p–value* < 0.05, * *p–value* < 0.1. [a] Robust s.e. in parentheses. *Flfp* stands for female labour force participation. *nmf* stands for children born out-of-wedlock (non-marital fertility). Columns 1 - 4 refer to dynamic FE (bias-corrected) and columns 5 - 9 refer to GMM difference.

On the rest of the independent variables, marriage is always negative, significant at the 1% level of significance (except for Table 3.2, column 6), and yields the greater impact on *nmf* among all the variables employed. Young women's labour participation has a negative sign but its significance varies. Female wage is also negative but there is no evidence of any significance. Lastly, female wage premium shows a positive effect and some statistical significance which is in accordance with Becker's marriage theory.

3.6 Conclusions

This paper tests the Easterlin relative income hypothesis in a non-marital fertility frame. Data from IPUMS-CPS, that spans a period of 35 years (1981-2016), has been employed to construct the relevant variables for the analysis. The goal is twofold. First, to examine for a potential correlation between relative income and non-marital fertility; and second, to find out which one of the absolute or relative income predicts better the outcome variable. Results obtained corroborate a negative and statistically significant effect of relative income on the fraction of children born out-of-wedlock. We also find that relative income exerts a greater impact than the absolute one.

Results are robust to the inclusion of marriage and other covariates, as well as to different econometric methods. On the former, the effect of relative income on births out-of-wedlock when marriage is considered, could probably indicate that young adults will act more carefully (i.e. use of contraception) and will avoid non-marital fertility, if they feel that they have odds to improve their socio-economic status. In other words, the gap between aspirations and income is not so large that can be filled. The latter is in line with evidence from the relevant literature that indicates cohabitation – and hence non-marital fertility – to prevail at lower socio-economic statuses (Kennedy and Bumpass, 2008; Manning and Cohen, 2015). A future work could examine whether the latter occurs especially for young individuals with lower relative income (those who miss the social ladder) rather than their peers of the same social class, who retain or improve their socio-economic profile.

Finally, the comparison showed that relative income emerges as a better predictor than the absolute one. Evidence provided in this paper reveals that young adults decide based rather on a relative perspective as Easterlin has claimed. From this point of view, relative income might be seen as a potential policy tool in order to reduce non-marital fertility and its repercussions.

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3.8 Appendix

A. Panel tests

Table A3.1 Pesaran (2007) unit-root test.

Variable	CPIS
Children out-of-wedlock	-4.691
Relative income	-4.738
Marriage	-4.441
Absolute income	-4.757
Female wage	-4.559
Flfp	-4.539
Male unemployment	-4.818
Premium	-4.684

Note: H_0: All panels contain unit-roots. Critical values: 10%:-2.05, 5%:-2.11, 1%:-2.23. *Flfp* stands for female labour force participation.

Table A3.2 Pesaran (2015) cross-section dependence test.

Variable	CD-test	*p*–value	Average joint T	Mean ρ	Mean abs (ρ)
Children out-of-wedlock	144.699	0.000	33.87	0.22	0.39
Relative income	43.194	0.000	33.87	0.22	0.34
Marriage	171.924	0.000	33.87	0.88	0.88
Absolute income	7.729	0.000	33.87	0.04	0.29
Female wage	11.085	0.000	33.87	0.06	0.30
Flfp	99.059	0.000	33.87	0.51	0.57
Male unemployment	62.443	0.000	33.87	0.32	0.40
Premium	17.816	0.000	33.87	0.09	0.38

Note: H_0: Cross-section independence. *Flfp* stands for female labour force participation.

Table A3.3 Wursten (2018) first order serial correlation test.

	HR-stat[a]	*p*–value
Residuals	0.66	0.511

Note: H_0: No first-order serial correlation. [a] Heteroskedasticity – robust Born and Breitung (2016) HR-test on residuals.

B. Data construction

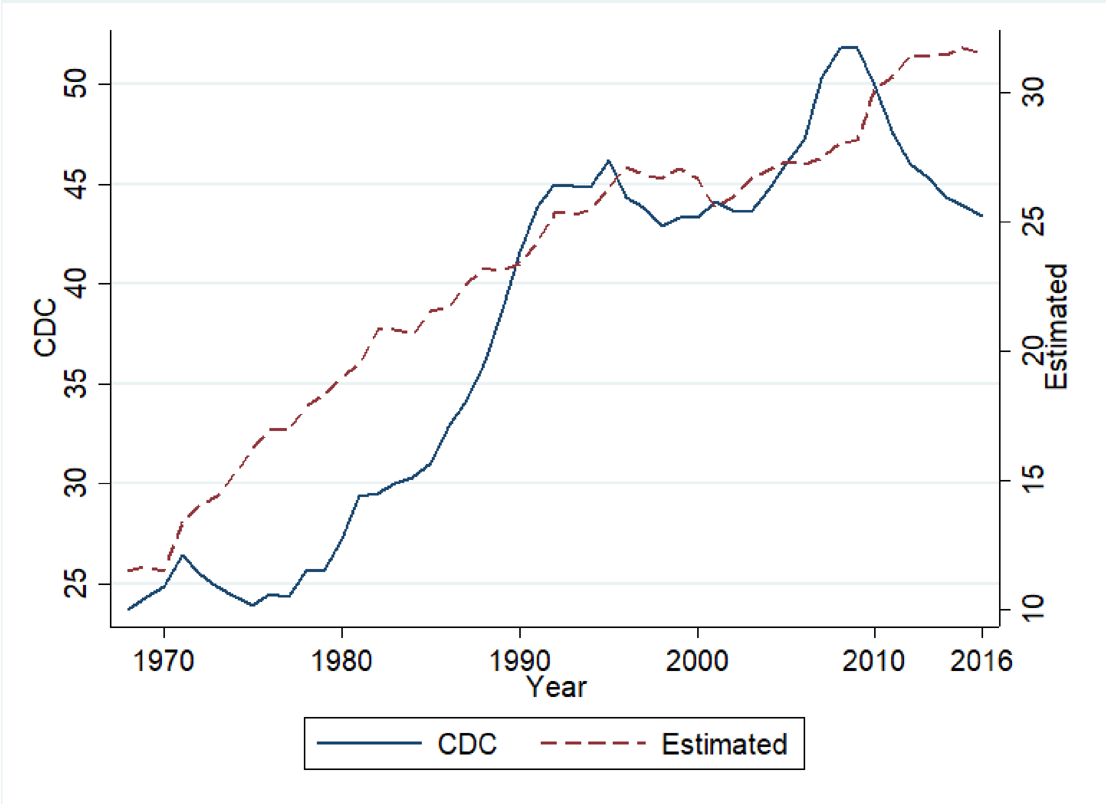

Fig. A3.1 Comparison for the fraction of children born out-of-wedlock for women aged 15-44 years old between data from CDC and our own calculations employing data from the IPUMS-CPS. *Source*: CDC and IPUMS-CPS.

Due to the lack of any disposal variable for non-marital fertility, we proxy the latter by calculating the fraction of the number of children out-of-wedlock by state and year in the US, for women aged 15-24 years old. We conduct this estimation by dividing the total number of women at this specific age group who are unmarried with own child/children in household, to the sum of the total number of women – independent of marriage – having child/children. To check the validity of the developed variable (children out-of-wedlock) to the non-marital fertility rate coming from the National Vital Statistics Reports (Fig. 3.1), we calculated the former once again restricted to the ages between 15 and 44. A strong correlation (0.916, Fig. A3.1) was found between the two variables. The latter provides our analysis with confidence over the developed variable as a proxy for non-marital fertility. Data constraints do not allow us to control directly for divorced, separated, or widowed women. Thus, we account for these effects by the use of decade dummies.

Fig. A3.2 depicts the evolution of relative income for white non-Hispanics aged 15-24 years old since 1981. We notice a decline until the middle of 90's followed by an increase and then stagnation around 2000. After a short decrease, we observe that relative income has been increasing since 2006.

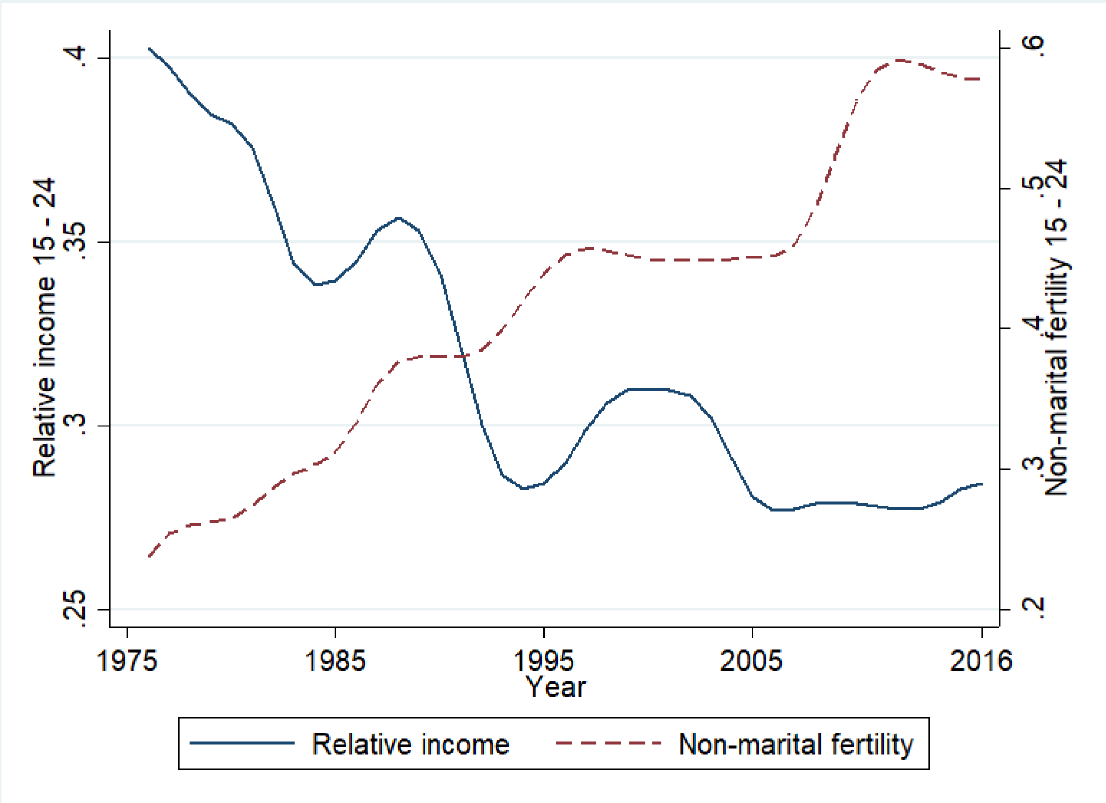

Fig. A3.2 Smoothed and normalized relative income of white non-Hispanics males aged 15-24 years old and children born out-of-wedlock (calculated) of the same age group. *Source*: Author's calculation employing data from IPUMS-CPS.

C. Control variables

We employ some control variables to address endogeneity problems and approach causal inferences. The absolute income is simply the numerator of relative income by which we compare it in the regression analysis. Except for these two incomes, there are also other factors one should consider of affecting children born out-of-wedlock. Consequently, we develop the following variables: female wage, female labour force participation, female wage premium, marriage, and male unemployment.

The use of dummies aims to capture time-varying effects which could affect non-marital fertility such as levels of inequality, poverty, gender ratio, contraception, divorce rates, etc. For the female wage development, we repeat the process followed for male absolute income. Female labour force participation rate was estimated as the percentage of women of a specific year-age-state who participate in the labour force to the total sum of women. In a similar way, unemployment rate was estimated as the fraction of unemployed men of the male labour force. Marriage rate is the ratio of individuals of both males and females – "marriage itself is a two-sex arrangement", Sweeney (2002:143) – who are married, to the sum of all women. Female wage premium is defined as the fraction of women's wages of age 25-65 holding at least a bachelor degree, to those of high school education of the same age group. All variables were estimated at a specific age-state-year level. Obviously, only female wage premium diverges of this specification. Variables were smoothed (running medians) and normalized to sum to 1 in order to address scale disparity. By normalizing, variables contribute equally to the analysis. In addition to this, we skip non-stationarity problems as the variables are bounded (Farmer, 2015).

Two year lags have been applied to the relative and absolute income aiming to capture potential reverse causality. It is reasonable to assume that both incomes would impact on non-marital fertility with a delay rather than straight forward; on average, nine months mediate between conception and birth day. Covariates have been lagged too. Labour participation of women could be affected by an increase of births out-of-wedlock due to the difficulty for a young unmarried woman to combine childrearing and work, or, childrearing and educational attainment – which would in turn increase odds to find a job. Male unemployment might be affected by reverse causality in the same way, but potentially less than the previous variables. The same mechanism might hold for female wages and female wage premium. We therefore impose lags to them too.

CHAPTER 4

Fertility Postponement, Female Education, and Young Men's Relative Deprivation

Abstract

This paper investigates the fertility postponement and the female educational attainment in the United States for the period 1981-2016. We consider Easterlin relative income hypothesis as a potential explanation. Easterlin claims that young men's standard of living is mostly formed at the parental household during childhood and competes to the one in young adulthood. This process is conceptualized by the notion of relative income. A lower level of the latter might lead young men to delay family formation. In turn, it affects young women who might response by increasing their educational attainment and labour force participation. If this process holds, fertility should delay too. We split this study into two parts. First, we test for a possible impact of relative income on the fertility postponement (*Hypothesis 1*). Second, we test for a possible impact of relative income on female education (*Hypothesis 2*). Both concepts are very closely related, given that literature has indicated female education as the main driver on later childbearing. Hence, *Hypothesis 1* could be perceived as a direct impact, whilst *Hypothesis 2* might be seen as an indirect one on the fertility postponement. Data retrieved from the IPUMS-CPS. Panel dynamic methods have been employed. Results show that relative income exerts a statistically significant effect on the fertility postponement (*Hypothesis 1*) as well as on the female educational attainment (*Hypothesis 2*).

Keywords: Easterlin hypothesis, fertility postponement, female education

JEL: J12, J13, I23

4.1 Introduction

A trend towards later childbearing has been taking place in the developed world during the last decades. The postponement is common in most of the European countries and the United States (US) since 70's. Fig. 4.1 depicts the latter in a sample of countries Europe and the US.

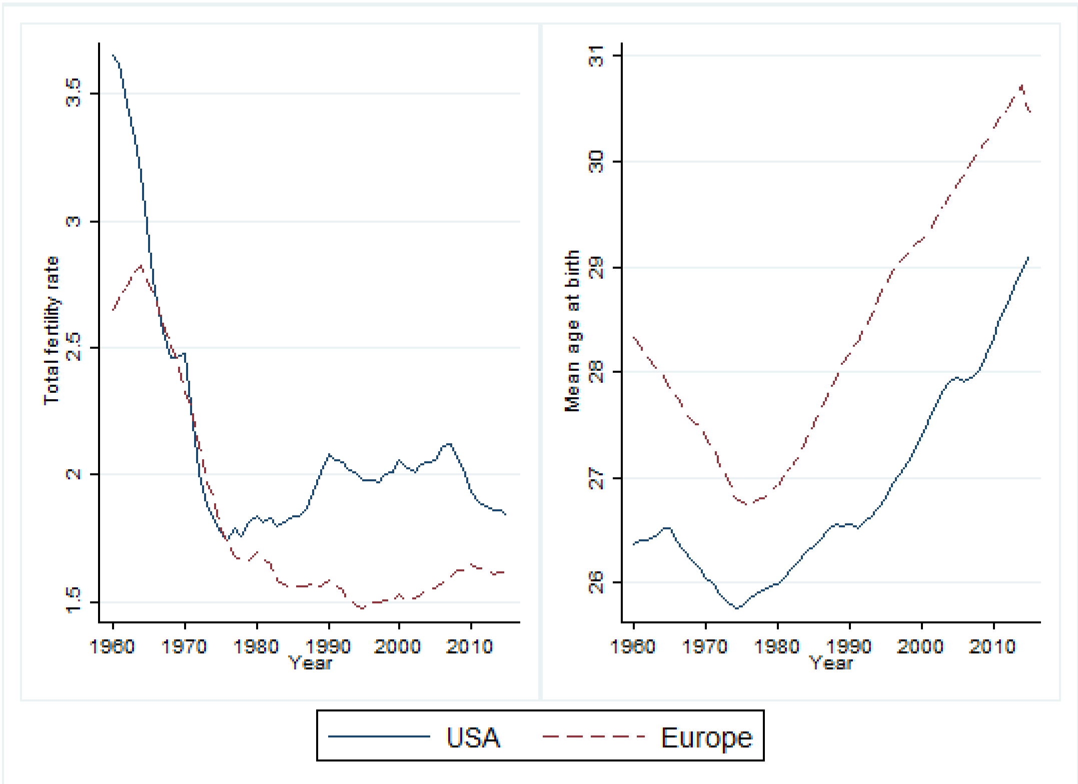

Fig. 4.1 Total fertility rate and mean age at birth of women. *Note***:** Europe refers to the mean of the corresponding index of the following European countries: Austria, France, Germany, Great Britain, Italy, Netherlands, and Switzerland. *Source*: author's calculation based on data extracted from the Human Fertility Database (HFD).

Research on the postponement determinants has become highly topical given the distortions observed to cause on the total fertility rate (TFR). To clarify, an increase of the mean age at birth results in period effects that obfuscates the true impact of potential pronatalist policies (Sobotka and Lutz, 2011; Bongaarts and Sobotka, 2012). In other words, as women postpone their fertility, TFR first declines and subsequently increases as women recuperate at older ages (a process known as tempo effect). Thus, it is not clear whether the observed uptrend after the decline is an outcome owing to pronatalist policies or to the postponement phenomenon. As a consequence, most demographers argue that the fertility rebound is an artifact of the latter.

A plethora of explanations has been provided with respect to the postponement. Women's labour force participation (Miller, 2011), the rising opportunity cost of childrearing (O'Donoghue et al., 2010), economic uncertainty (Goldin and Katz, 2002) and cultural shifts (Beck and Beck-Gernsheim, 2001), are some of the examined factors (for a review, see Mills et al., 2011). Among all, female education has been indicated as the most profound (Ní Bhrolcháin and Beaujouan, 2012). During the last decades, except for the fertility postponement, women have been increasing their educational attainment too. While the relationship between female education and fertility postponement has been widely investigated (Rindfuss et al., 1980; Lappegard and Ronsen, 2005; Neels et al., 2017), less is known on the causes of the increased female education (Norton and Tomal, 2009). The investigation of the latter would potentially shed more light on the fertility postponement too, given their highly correlation.

This study draws data from the IPUMS-CPS for the period 1981-2016 across the US. We employ dynamic panel data methods to investigate whether the relative deprivation of young men (Easterlin, 1966) could explicate the female education and the mean age at first birth. Two cases are probed. The first one considers a direct impact on the mean age at first birth. The second one takes advantage of the fact that women's education has been suggested as the main driver of the fertility postponement. Hence, providing evidence on the relationship between relative income and female education, one could perceive it as an extra indication towards an indirect impact (through female education) on the fertility postponement. Hence, we aim to investigate the following two hypotheses:

Hypothesis 1: the relative income decline of young men coincides with shifts to later childbearing (direct impact on the fertility postponement).

Hypothesis 2: the relative income (RY henceforth) decline of young men coincides with increases to female educational attainment (also seen as an indirect impact on the fertility postponement).

To our knowledge, this paper is the first to attribute fertility postponement to young men's relative deprivation. This is also the first paper to relate young men's relative deprivation to young women's educational attainment. In this way, we add on both the demographic and the female educational literature. Lastly, given the demographic aspect for an artifact fertility rebound caused by the tempo effect, we conclude this paper with some thoughts on the potential contribution of RY to the recent fertility rebound.

4.2 Background

4.2.1 Literature review

The relationship between education and age at parenthood has been well-established in the literature. Higher educated women usually postpone their decision on motherhood to later ages (Rindfuss et al., 1988, 1996; Martin, 2000). Many mechanisms have been indicated for the latter. Female education increases career aspirations which subsequently delays family formation (Lappegård and Rønsen, 2005), because being enrolled as a student and the role of mother is difficult to be reconciled. The higher opportunity cost of the career-oriented women would further reinforce family postponement. In addition, education might also lead to more individualistic values (van de Kaa, 1987) and hence on fertility delays. However, other scholars question the proposition that rising educational attainment leads to later family formation (see Oppenheimer et al., 1995 and Rendall et al., 2010).

On the other hand, the discussion of the causes on the female educational attainment is rather scarce in the literature. One could assume that the fertility fall allowed girls to spend more time in school rather than taking care of their siblings.[1] At the same time, increasing family incomes make parents more capable to spend resources on girls' education too (Heath and Jayachandran, 2016). Another explanation states that fertility decline permit women to participate in the labour force.[2] In turn, women proceed to higher education to increase their wages (Macunovich, 2003). Women who target on higher earnings (career-oriented) engage in activities to acquire more human capital. However, it could be the other way round: female wage increment attracted more women to participate in the labour market and subsequently to delay fertility. Thus, one ends up to a causal ""chicken or egg" problem" (Macunovich, 2003: 109).

4.2.2 A potential path from relative income to female education and fertility postponement

The relative deprivation of young men might be the common factor behind female education and fertility postponement. One plausible path could be related to marriage rates. The decline of young men's relative income results in marriage retreat due to the increasing gap between incomes and material aspirations (Macunovich, 2011). Thus, young men might be hesitated in family formation if they feel that they lack the desired standard of living. Accordingly, they will proceed to marriage postponement and hence odds to an early marriage will worsen for women. Seen strictly from this perspective (men's marriage postponement decision), one could argue that women might not choose to delay family formation rather they are "enforced" by young men's behaviour.

Becker (1991) argues that marriage turns up some similarities with the international trade in the sense that men and women specialize in a specific domain; men in labour market and women in household work. In this context, women's "comparative advantage" within marriage is set aside in young adulthood, if men decide to delay (or even forgo) marriage at later stages of their

1 Notice that this explanation presumes that fertility fall is the cause and educational rise the outcome. This paper argues and tries to demonstrate the opposite.

2 This argument implies that fertility decline is exogenous to women's education and labour force participation.

lifes. Accordingly, women engage in activities to enhance their position in the society. Higher educational attainment could be perceived as the most relevant act for the latter. However, and in any case, since the period between leaving school and time of marriage widens, women engage in other activities instead of waiting for their marriage.[3] Thus, men's marriage delay may have further contributed to the empowerment of women. The aforementioned path might be thought as one of the reasons on the increasing female labour force participation and the educational attainment during the last decades in the US (Greenwood et al., 2016). Hence, one could infer that woman's higher levels of educational attainment – and subsequent fertility postponement – could be attributed to the feelings of relative deprivation among young men who decide to delay their marriage. Of course, one could not attribute such complex social phenomena solely to a factor (young men's RY). RY might be thought as one of the many potential reasons lying behind it.

4.3 Data description

We employ data from the IPUMS-CPS database which spans the period 1981-2016 for the US. The panel considered is unbalanced. The fixed-effect (FE) estimator is still consistent under unbalanced panels as long as missing observations are random (Wooldridge, 2002). A drawback comes from the loss of observation (less power of the test). An attempt to reshape the unbalanced panels into balanced (as we are forced to do for the causality test in 4.5.3) would result in an even greater loss of observations. Thus, we leave our panel

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[3] A reasonable critique to this aspect could be the increase of cohabitation against marriage. However, cohabitation is not a perfect substitute to marriage. The former appears mostly as a precursor on the latter (Bumpass and Lu, 2000; Xie et al., 2003). Thus, evidence weakens such a critique. We try to deal with cohabitation by the inclusion of decade dummies. Unfortunately, IPUMS-CPS provides data on cohabitation since 2007. Note also that decade dummies could also control for the social changes that took place during the years under examination which facilitated the entrance of women in higher education. Therefore, they additionally aim to capture the social shifts occurred in life pattern through years.

unbalanced since we come under the case where missing observations among states is random.

The annual social and economic supplement (*asec*) of the IPUMS database has been used to construct the RY for young men and the percentage of higher educated young women. The analysis takes place in a state-year-age level basis. The focus of the study is on white non-Hispanics due to their greater representativeness in the database. Moreover, we keep in the analysis only those who are civilians as in Macunovich (2012).

RY has been developed as a fraction. The numerator consists of the median wages for men aged 15-24 years old (full time-year workers). We also dropped individuals from the analysis if their estimated wage per hour was less than \$2.50 or more than \$250 in 2008 constant dollars (Blau and Kahn, 2007) to confront heterogeneity. The denominator denotes the median family income with a "Head" of either sex aged 45-54 years old. The latter aims to instrument aspirations. In contrast to the numerator, here, we do not constraint to any race or income level. We take into account all individuals of each state without exclusions (see p. 43).

We set the threshold of high school for the female educational attainment construction. The age at which students end high school in the US is roughly 18 years old. Therefore, the minimum threshold for the calculation of the educational percentage is the latter. Female education has been calculated by dividing the number of women aged 18-24 years old who have attained equal or more than high school, to the sum of all women of the same age group. Mean age at first birth has been calculated as the mean age of women when their first child was born by year, state, and race. The fertility supplement of the IPUMS-CPS provides the necessary information. The latter led to a great loss of observations due to its highly restriction; the age at first birth is offered for the following years: 1977, 1979, 1980-1983, 1986-1988, 1990, 1992, 1995, 2012, 2014, and 2016. Since the fertility supplement does not provide direct data on the age of women at first birth, we calculated it applying the following formula: age at 1st birth = year of 1st birth – (current year – current age).

Finally, we have included the female labour force participation, male unemployment, and male marriage rate as control variables. Decade dummies have been employed to control for time-varying unobserved heterogeneity (e.g., shift in ideologies, social norms, cohabitation, etc.). Variables have been smoothed (based on running medians) and normalized to sum to unity. In this way we address scale disparity and non-stationarity problems as they become bounded (Farmer, 2015). Normalizing process has been set in order to equalize the contribution of the variables. Two year lags have been applied on the independent variables to address potential reverse causality.

4.4 Methodology

We employ: (1) fixed-effects (FE) and bias-corrected FE (Bruno, 2005a, 2005b); (2) two stage least squares FE (2SLS FE, see Schaffer, 2015) and GMM difference (Arellano and Bond, 1991; Roodman, 2009); (3) Granger non-causality tests (Dumitrescu and Hurlin, 2012). The FE estimator controls for time-constant variables that have not been incorporated into the model. We include decade dummies to control further for any time-varying effects. 2SLS and GMM estimators have been employed to confront endogeneity. In the absence of any other accurate instrument, the lagged value of relative income has been employed. Lastly, we control for autocorrelation, heteroscedasticity, and cross-sectional dependence. The first model to be estimated is as follows:

$$MAB_{i,t} = b_0 + b_1 RY_{i,t-2} + \sum_{j=2}^{4} b_j X'_{i,t-2} + \sum_{i=1}^{4} b_{5,i} dummy_i + \mu_i + u_{i,t} \quad (1)$$

where $MAB_{i,t}$ denotes the mean age at first birth, $RY_{i,t-2}$ is the relative income (with a two years lag), $X'_{i,t-2}$ is a vector of control variables (with a two years lag), $dummy_i$ takes the value 1 for the corresponding decade and 0 otherwise, μ_i is the fixed-effect term, and $u_{i,t}$ is the stochastic error term.

Next, we consider the following model to test for the female education:

$$Education(f)_{i,t} = b_0 + b_1 Education(f)_{i,t-1} + b_2 RY_{i,t-2} + \sum_{j=3}^{5} b_j X'_{i,t-2} +$$
$$+ \sum_{i=1}^{3} b_{6,i} dummy_i + \mu_i + u_{i,t} \qquad (2)$$

where $Education(f)_{i,t}$ denotes the percentage of women attained equal or more than high school, $RY_{i,t-2}$ is the relative income (with a two years lag), $X'_{i,t-2}$ is a vector of control variables (with a two years lag), $dummy_i$ takes the value 1 for the corresponding decade and 0 otherwise, μ_i is the fixed-effect term, and $u_{i,t}$ is the stochastic error term.

We apply the Granger non-causality procedure proposed by Dumitrescu and Hurlin (2012) to test for any causal effects on women's educational attainment. Although this method accounts for heterogeneous panel data models (Dumitrescu and Hurlin, 2012: 22), we transform our panel into balanced in order to carry out the test (see Lopez and Weber, 2017: 973). The cost is a drop in observations (1650 to 1024). The specification can be written as:

$$Education(f)_{i,t} = a_i + \sum_{k=1}^{K} \beta_i^{(k)} Education(f)_{i,t-k} + \sum_{k=1}^{K} \gamma_i^{(k)} RY_{i,t-k} + \varepsilon_{i,t} \qquad (3)$$

where k is the number of lags. The null hypothesis (no causality) is defined as:

$$H_o\colon \gamma_{i1} = \ldots = \gamma_{ik} = 0 \qquad \forall\ i = 1, \ldots, N$$

and the alternative hypothesis (causality) as:

$$H_1\colon \gamma_{i1} = \ldots = \gamma_{ik} = 0 \qquad \forall\ i = 1, \ldots, N_1$$
$$\gamma_{i1} \neq 0\ or \ldots or\ \gamma_{ik} \neq 0 \qquad \forall\ i = N_1 + 1, \ldots, N$$

where $N_1 \in [0, N-1]$ is unknown. This test is based on the individual Wald statistics of Granger non-causality averaged across the cross section units:

$$\bar{W} = \frac{1}{N} \sum_{i=1}^{N} W_i$$

W_i is the standard adjusted Wald statistic for individual i observed during T periods. This test accounts for a possible difference of the unconstrained parameters from one individual (state) to another. Moreover, it allows for two subgroups of cross-section units and cross-sectional dependence too (Dumitrescu and Hurlin, 2012: 3-4). It should be emphasized that rejecting the null hypothesis does not exclude the possibility of no causality for some individuals (states).

The number of lags has been selected based on Bayesian information criterion. Variables must be stationary (Lopez and Weber, 2017: 5). The Pesaran (2007) unit-root test for panel data ensures the latter. In all cases the null hypothesis of non-stationarity was rejected. Results are available upon request.

4.5 Empirical results

Fig. 4.2 plots the RY with respect to female education and mean age at first birth. The negative correlation that emerges for both relationships is well-fitted. Next, we keep on with the presentation of the obtained regression results.

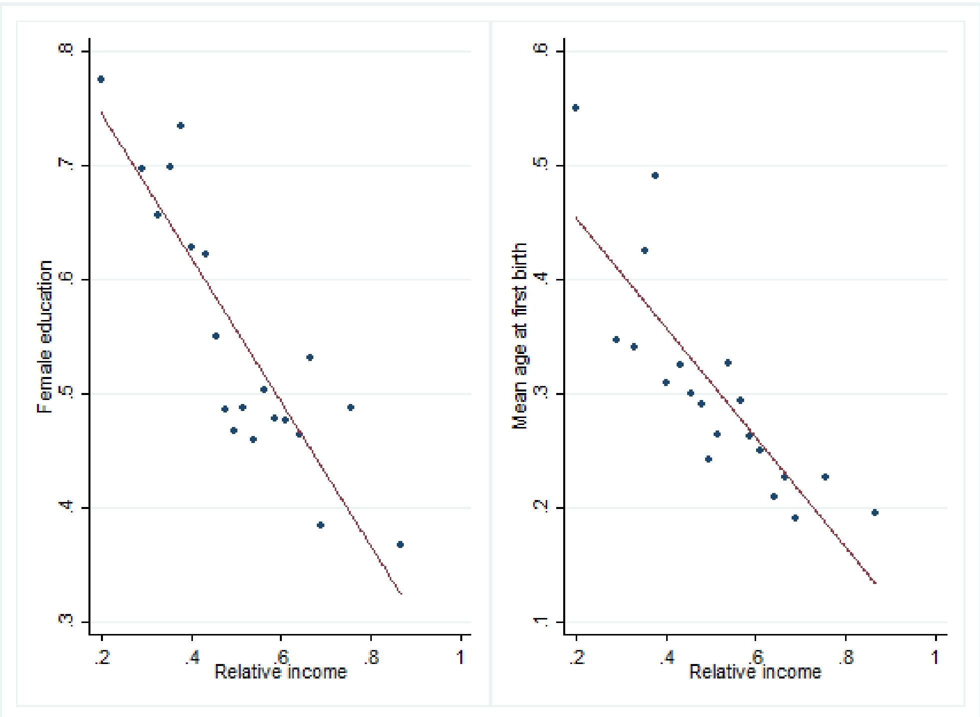

Fig. 4.2 Binscatter plots. *Note*: plots on the relationship between relative income – female education, and relative income – mean age at first birth. *Source*: Author's calculation based on data extracted from the IPUMS-CPS.

4.5.1 Relative income and fertility postponement (*Hypothesis 1*)

First, we examine the hypothesis that RY is a significant contributor to the fertility postponement (*Hypothesis 1*). In this case, we follow a non-dynamic approach in order to avoid further reduction to the number of observations which is already limited. In Table 4.1, we observe that RY is negative and statistically significant in all cases examined. The degree of the significance varies with respect to the method employed. For the FE, it ends up to the 10% level of significance when control variables and decade dummies are added (column 3).

Table 4.1 Regression results for mean age at 1st birth (parental age 45-54).

Dependent variable:	**FE**			**2SLS FE**		
Mean age at 1st birth	**(1)**	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**
RY_{t-2}	-0.549***	-0.179*	-0.169*	-0.490***	-0.262***	-0.253***
	(0.0972)	(0.0933)	(0.0990)	(0.0574)	(0.0845)	(0.0879)
$Education(f)_{t-2}$		0.307***	0.218**		0.185***	0.176***
		(0.0576)	(0.0936)		(0.0459)	(0.0494)
$Flfp_{t-2}$		0.0551	0.105		0.0684	0.0714
		(0.110)	(0.118)		(0.0926)	(0.0928)
$Unemployment(m)_{t-2}$		0.0477	0.0475		0.0140	0.0121
		(0.105)	(0.106)		(0.0401)	(0.0388)
$Marriage(m)_{t-2}$		-0.118	-0.105		-0.0561	-0.0562
		(0.0809)	(0.0852)		(0.0598)	(0.0591)
Decade dummies	No	No	Yes	No	No	Yes
Constant	0.606***	0.246**	0.232**	0.606***	0.246**	0.232**
	(0.0495)	(0.108)	(0.114)	(0.0495)	(0.108)	(0.114)
R-squared	0.280	0.515	0.521	0.445	0.722	0.725
Observations	244	244	244	137	137	137

Note: Robust standard errors in parentheses. *** *p–value* < 0.01, ** *p–value* < 0.05, * *p–value* < 0.1. Columns 1-3 refer to FE estimator and 4-6 to 2SLS FE estimator (instrumental variable: two years lag of RY); *m* denotes males and *f* females.

Controlling for endogeneity (columns 4-6), we see that RY keeps its negative sign. Significance still remains at the 1% level after the inclusion of the control variables and decade dummies (columns 5 and 6, respectively). Thus, Table 4.1 shows that mean age at first birth is negatively affected by the RY. Consistent with prior research, women's education presents a positive and statistically significant effect in all columns. Female labour force participation, as well as men's unemployment rate, found positive, albeit insignificant. Men's marriage rates are negative and insignificant too.

It is important to notice that according to Table 4.1, RY displays an "independent effect" on the mean age at first birth. By independent effect we imply that significance remains after controlling for women's education, which has been considered as the main mediator (Table 4.2 also confirms). This finding may also suggest the presence of other paths that have not been considered.

4.5.2 Relative income and female education (*Hypothesis 2*)

This subsection examines whether young men's RY coincides with higher educational attainment of women (*Hypothesis 2*). Here, we also take advantage of the fact that women's education has been suggested as the main driver on later childbearing so we can infer for the latter too.

Table 4.2 Regression results for female education (parental age 45-54).

Dependent variable:	FE (bias-corrected)			GMM difference		
Education(f)	(1)	(2)	(3)	(4)	(5)	(6)
Education(f)$_{t-1}$	0.909***	0.871***	0.647***	0.273***	0.382***	0.355***
	(0.0082)	(0.0116)	(0.014)	(0.0803)	(0.0828)	(0.0828)
RY$_{t-2}$	-0.148***	-0.157***	-0.063***	-0.362***	-0.343**	-0.461***
	(0.018)	(0.018)	(0.015)	(0.108)	(0.136)	(0.149)
Flfp$_{t-2}$		0.083***	0.001		0.243**	0.194*
		(0.017)	(0.015)		(0.103)	(0.112)
Unemployment(m)$_{t-2}$		-0.069***	-0.047***		0.0248	0.212***
		(0.015)	(0.012)		(0.0613)	(0.0727)
Marriage(m)$_{t-2}$		-0.108***	-0.073***		0.246***	0.215***
		(0.017)	(0.014)		(0.0577)	(0.0606)
Decade dummies	No	No	Yes	No	No	Yes
Number of instruments				47	47	47
Observations	1,583	1,583	1,583	1,442	1,442	1,442
F-stat.				0.000	0.000	0.000

Note: Robust standard errors in parentheses. *** *p-value* < 0.01, ** *p-value* < 0.05, * *p–value* < 0.1. Columns 1-3 refer to FE bias-corrected estimator and 4-6 to GMM difference estimator (instrumental variable: two years lag of RY); *m* denotes males and *f* females.

In Table 4.2, we see that relative income is always negative and statistically significant (1% level). This holds for both methods. However, results on the rest of the variables are qualitatively different. Men's unemployment and marriage rates switch from negative (bias-corrected FE) to positive (GMM difference). In this case, we rely on the results obtained by the bias-corrected FE estimator. The small number of T (=35) might cause overfitting in variables instrumented by GMM. Bias-corrected dynamic panel estimator is more appropriate for such a moderate values of T in panel data analysis (for more on this, see Bruno, 2005a; 2005b).

The negative sign on men's unemployment might be explained by the decreasing returns to human capital when unemployment rate is high. We observe that RY remains significant after controlling for male marriage rates. Therefore, although results confirm the suggested path (through male marriage), they additionally indicate other path(s) that could be in place too.

4.5.3 Causality tests

Finally, we test for any causal effects between relative income and female educational attainment. We are not able to test for a causal effect with respect to mean age at first birth due to the low number of observations. On the former, any causal finding might be perceived as an evidence for a potential contribution on the mean age at first birth, as explained above. We apply the Dumitrescu and Hurlin (2012) Granger non-causality test (Lopez and Weber, 2017).[4] The lag order for the implementation of the test has been selected according to the Bayesian Information Criterion (BIC). The test has been conducted in both directions (RY → female education; female education → RY).

Results are presented in Tables 4.3 and 4.4. In Table 4.3, the *p-value* equals to zero rejecting the null hypothesis of no causality at the 1% level of significance (at least). Hence, we can conclude that there is a causal effect of

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[4] Data transformed into a balanced panel. We have already shown that the necessary condition for this transformation holds: variables are stationary.

relative income on female education for at least one state. On the reverse direction (Table 4.4), *p-value* implies that the null hypothesis of no-causality can be rejected too, but with less significance (0.051).

Table 4.3 Dumitrescu and Hurlin (2012) Granger non-causality test.

Statistic	Value	Critical value
$\bar{W}$	52.1611	
$\bar{Z}$	62.4532	Not obtained
p-value	(0.0000)	
$\tilde{Z}$	19.8768	Not obtained
p-value	(0.0000)	

H_0: Relative income does not Granger-cause female education.
H_1: Relative income does Granger-cause female education for at least one state.

Table 4.4 Dumitrescu and Hurlin (2012) Granger non-causality test.

Statistic	Value	Critical value
$\bar{W}$	27.8792	
$\bar{Z}$	28.1134	28.1918
p-value	(0.0510)	
$\tilde{Z}$	8.0938	8.1207
p-value	(0.0510)	

H_0: Female education does not Granger-cause relative income.
H_1: Female education does Granger-cause relative income for at least one state.

Note for Tables 4.3 and 4.4: Lag order has been selected based on the Bayesian information criterion (BIC). $\bar{W}$: Wald statistic, $\bar{Z}$: standardized statistic, $\tilde{Z}$: standardized statistic for large N but relatively small T datasets. *p-value* in parentheses computed using 1000 bootstrap replications.

In Appendix (Tables C4.1 and C4.2), where we present the robustness test by changing the parental age group, results are mixed. Thus, we see that the null hypothesis of no causality can be rejected for the path that leads from RY to female education in both age groups considered. On the other way round, the null is rejected only for the 50-59 age group. Thus, the Granger non-causality test inclines to a causal effect from RY to female education rather the inverse.

4.5.4 Robustness check

By the selection of the specific age group of young adults (15-24) and the assumed mean age at birth (30) of their parents[5], we conclude that the parental age will range between 45 and 54 (15-24, +30) years old, which we lag by 5 years. We have adopted the age of thirty at the core analysis, since we aim to capture aspirations for all possible birth orders of young adults in the sample, and at the same time, for different educational levels of their parents (either male or female). As a robustness check, we use two more parental age groups in the denominator of the relative income to control for differences in mean age at birth of young men's parents. Hence, mean ages at birth of 25 and 35 years old have been also regarded. The corresponding age groups with respect to the family income in childhood are 40-49 and 50-59, respectively. To sum up, we have tested for three different parental birth age groups: 40-49 for mean age at birth=25, 45-54 for mean age at birth=30, and 50-59 for mean age at birth=35. Results are presented in Appendix A, B, and C. They are qualitatively similar to the ones in the main text.

4.6 Conclusions

This study uses data from the IPUMS-CPS database that captures a period of 36 years (1981-2016) in the US. We developed a measure of Easterlin's relative income. We tested the hypothesis that the drop of young men's RY in the US coincides to increments in mean age at first birth of women (*Hypothesis 1*). Conventional wisdom claims that the most important factor on later childbearing is the female educational attainment. Thus, we further examined the hypothesis that the fall of relative income coincides with increases to female educational attainment (*Hypothesis 2*). Both hypotheses have been found statistically significant and signs have been the ones expected.

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[5] For male's mean age at birth in US, see Khandwala et al. (2017), and for female's, see Mathews et al. (2002, 2016).

Next, we examined the possibility of a Granger causal relationship, instead of simple correlation only for the case of female education and RY. The low number of observations did not allow us to test for causality in *Hypothesis 2* (fertility postponement). The direction found to run from RY to female education rather than the opposite. This opposite direction potentially might be explicated by the subsequent increase of the female labour force participation which suppressed young men's wages and – probably – increased their material aspirations through the increased family income in childhood.

The potential paths assumed in section 4.2 are supported by our analysis. Obtained results also suggest that fertility postponement may be the result of more mediators than the ones suggested (female education and male marriage). We reveal the effects of the two latter and indicate other(s) for future investigation because: (1) RY found significant on female education after controlling for male marriage and, (2) RY found significant on mean age at first birth after controlling for female education and male marriage.

Finally, the path examined in this paper could be also extended in the context of the fertility rebound (Myrskyla et al., 2009). Literature suggests many factors on the latter (for a review, see Balbo et al., 2013). Beyond other socio-economic arguments (Myrskyla et al., 2009; Luci-Greulich and Thevenon, 2014; Anderson and Kohler, 2015; Yakita, 2018), demographers put forward the tempo effect approach. However, our findings might imply that the tempo effect takes place under the influence of socio-economic aspects that have not been considered before. Accordingly, one could argue that the gap between incomes and material aspirations of young men has a significant effect on female education, and subsequently on fertility delay. From this point of view, what demographers have called an "artifact" of the fertility rebound might be rooted back to socio-economic reasons related to young men, instead of women's education or/and their labour force participation itself. This is important because particularly the latter has been indicated as a causal factor on the fertility evolution (Salamaliki et al., 2013). A future investigation within the European frame might shed more light on a potential linkage between Easterlin's RY and the fertility rebound.

4.7 References

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4.8 Appendix

A. Relative income and mean age at 1st birth

Table A4.1 Regression results for mean age at first birth (parental age: 40-49).

Dependent variable:	**FE**			**2SLS FE**		
Mean age at 1st birth	**(1)**	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**
RY_{t-2}	-0.592***	-0.251**	-0.242**	-0.502***	-0.270***	-0.274***
	(0.0949)	(0.0965)	(0.100)	(0.0468)	(0.0867)	(0.0919)
$Education(f)_{t-2}$		0.281***	0.207**		0.180***	0.183***
		(0.0596)	(0.0960)		(0.0425)	(0.0454)
$Flfp_{t-2}$		0.0610	0.106		0.0347	0.0333
		(0.104)	(0.114)		(0.0746)	(0.0770)
$Unemployment(m)_{t-2}$		0.0319	0.0313		0.00450	0.00505
		(0.101)	(0.105)		(0.0378)	(0.0373)
$Marriage(m)_{t-2}$		-0.0979	-0.0886		-0.0160	-0.0154
		(0.0832)	(0.0888)		(0.0600)	(0.0598)
Decade dummies	No	No	Yes	No	No	Yes
Constant	0.635***	0.294**	0.281**			
	(0.0495)	(0.116)	(0.123)			
R-squared	0.362	0.532	0.537	0.582	0.752	0.752
Observations	244	244	244	137	137	137

Table A4.2 Regression results for mean age at first birth (parental age: 50-59).

Dependent variable:	**FE**			**2SLS FE**		
Mean age at 1st birth	**(1)**	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**
RY_{t-2}	-0.429***	-0.0749	-0.0597	-0.388***	-0.153**	-0.142*
	(0.0836)	(0.0843)	(0.0924)	(0.0468)	(0.0780)	(0.0805)
$Education(f)_{t-2}$		0.328***	0.216**		0.208***	0.191***
		(0.0588)	(0.0944)		(0.0467)	(0.0500)
$Flfp_{t-2}$		0.00751	0.0647		-0.00204	0.00804
		(0.112)	(0.122)		(0.0824)	(0.0859)
$Unemployment(m)_{t-2}$		0.0400	0.0446		-0.00776	-0.00972
		(0.106)	(0.107)		(0.0402)	(0.0389)
$Marriage(m)_{t-2}$		-0.108	-0.0920		-0.0322	-0.0345
		(0.0856)	(0.0895)		(0.0732)	(0.0720)
Decade dummies	No	No	Yes	No	No	Yes
Constant	0.524***	0.207*	0.189			
	(0.0385)	(0.110)	(0.115)			
R-squared	0.255	0.502	0.509	0.385	0.695	0.699
Observations	244	244	244	137	137	137

Note for Tables 4.A1, 4.A2, 4.B1, 4.B2: Robust standard errors in parentheses, *** *p-value* < 0.01, ** *p-value* < 0.05, * *p-value* < 0.1. Columns 1-3 refer to FE (or bias corrected) estimator and 4-6 to 2SLS FE (or GMM difference) estimator (Instrumental variable: two years lag of RY). *m* denotes males and *f* females.

B. Relative income and female educational attainment

Table B4.1 Regression results for female education (parental age group: 40-49).

Dependent variable:	**FE (bias corrected)**			**GMM difference**		
Education(f)	**(1)**	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**
Education(f)$_{t-1}$	0.923***	0.877***	0.651***	0.393***	0.469***	0.446***
	(0.008)	(0.011)	(0.012)	(0.0534)	(0.0637)	(0.0641)
RY$_{t-2}$	-0.082***	-0.101***	-0.022**	-0.108**	-0.139*	-0.227**
	(0.014)	(0.014)	(0.011)	(0.0444)	(0.0774)	(0.0909)
Flfp$_{t-2}$		0.085***	-0.003		0.258***	0.193**
		(0.018)	(0.015)		(0.0876)	(0.0901)
Unemployment(m)$_{t-2}$		-0.116***	-0.077***		0.000171	0.179**
		(0.017)	(0.014)		(0.0522)	(0.0700)
Marriage(m)$_{t-2}$		-0.075***	-0.046***		0.285***	0.262***
		(0.015)	(0.012)		(0.0625)	(0.0651)
Decade dummies	No	No	Yes	No	No	Yes
Constant						
Number of instruments				47	47	47
Observations	1,583	1,583	1,583	1,442	1,442	1,442
F-stat.				0.000	0.000	0.000

Table B4.2 Regression results for female education (parental age group: 50-59).

Dependent variable:	**FE (bias corrected)**			**GMM difference**		
Education(f)	**(1)**	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**
Education(f)$_{t-1}$	0.906***	0.872***	0.6449***	0.291***	0.355***	0.335***
	(0.009)	(0.011)	(0.012)	(0.0709)	(0.0852)	(0.0888)
RY$_{t-2}$	-0.125***	-0.131***	-0.064***	-0.326***	-0.459**	-0.586***
	(0.017)	(0.017)	(0.013)	(0.0816)	(0.172)	(0.193)
Flfp$_{t-2}$		0.083***	-0.0002		0.260**	0.271*
		(0.018)	(0.0151)		(0.118)	(0.143)
Unemployment(m)$_{t-2}$		-0.100***	-0.048***		0.0947	0.291**
		(0.017)	(0.0120)		(0.0853)	(0.111)
Marriage(m)$_{t-2}$		-0.069***	-0.0478***		0.243***	0.216**
		(0.015)	(0.0120)		(0.0856)	(0.0954)
Decade dummies	No	No	Yes	No	No	Yes
Constant						
Number of instruments				47	47	47
Observations	1,583	1,583	1,583	1,442	1,442	1,442
F-stat.				0.000	0.000	0.000

C. Granger non-Causality tests

Table C4.1 Dumitrescu and Hurlin (2012) Granger non-causality test.

Parental age group	**40-49**		**50-59**	
Statistic	**Value**	**Critical value**	**Value**	**Critical value**
$\bar{W}$	46.2477		43.2579	
$\bar{Z}$	54.0904		49.8622	
p-value	(0.0000)		(0.0000)	
$\tilde{Z}$	17.0073		15.5565	
p-value	(0.0000)		(0.0000)	

H_0: Relative income does not Granger-cause female education.
H_1: Relative income does Granger-cause female education for at least one state.

Table C4.2 Dumitrescu and Hurlin (2012) Granger non-causality test.

Parental age group	**40-49**		**50-59**	
Statistic	**Value**	**Critical value**	**Value**	**Critical value**
$\bar{W}$	22.5819		24.6190	
$\bar{Z}$	20.6219	28.8141	23.5028	31.2790
p-value	(0.1950)		(0.0010)	
$\tilde{Z}$	5.5232	8.4088	6.5117	9.1800
p-value	(0.1950)		(0.0010)	

H_0: Female education does not Granger-cause relative income.
H_1: Female education does Granger-cause relative income for at least one state.

Note for Tables C4.1 and C4.2: *p-value* in parentheses computed using 1000 bootstrap replications. Lag order has been selected based on the Bayesian information criterion (BIC). $\bar{W}$: Wald statistic, $\bar{Z}$: standardized statistic, $\tilde{Z}$: standardized statistic for large N but relatively small T datasets.

CHAPTER 5

On the Drivers of the Fertility Rebound

Abstract

This paper investigates empirically the fertility rebound in low and high income OECD countries for the period 1970 to 2016. The focus is on the turning points of the rebound for the two country groups. We estimate the turning points in terms of GDP per capita, labour productivity, and female labour force participation. Results suggest that the rebound, (1) is statistically significant for low and high income OECD countries in terms of GDP per capita and labour productivity, (2) is present for female labour force participation mostly for the higher income country group and, (3) differences in the turning points are minimized for the two country groups when labour productivity is considered. Labour productivity emerges as the most important economic driver for the fertility rebound.

Keywords: fertility rebound, economic development, labour productivity

JEL Classification Codes: J11, J13, J21

5.1 Introduction

In late 90's – early 00's many developed countries experienced an increase in their total fertility rates (TFR) for the first time after decades of continuously drops. The first as well as the second demographic transition (SDT) (Lesthaeghe, 2010) refer to a steady decline in fertility rates along with increases in human (economic) development. However, the well-established fact of a negative association between fertility and economic development (Kirk, 1996; Doepke, 2004) has been questioned. Myrskylä et al. (2009) employing cross-section and longitudinal analyses shows that higher levels of socio-economic development coincide with fertility increments. Hence, a new pattern arises which changes the relationship between development and fertility from negative to an inverse J-shaped.

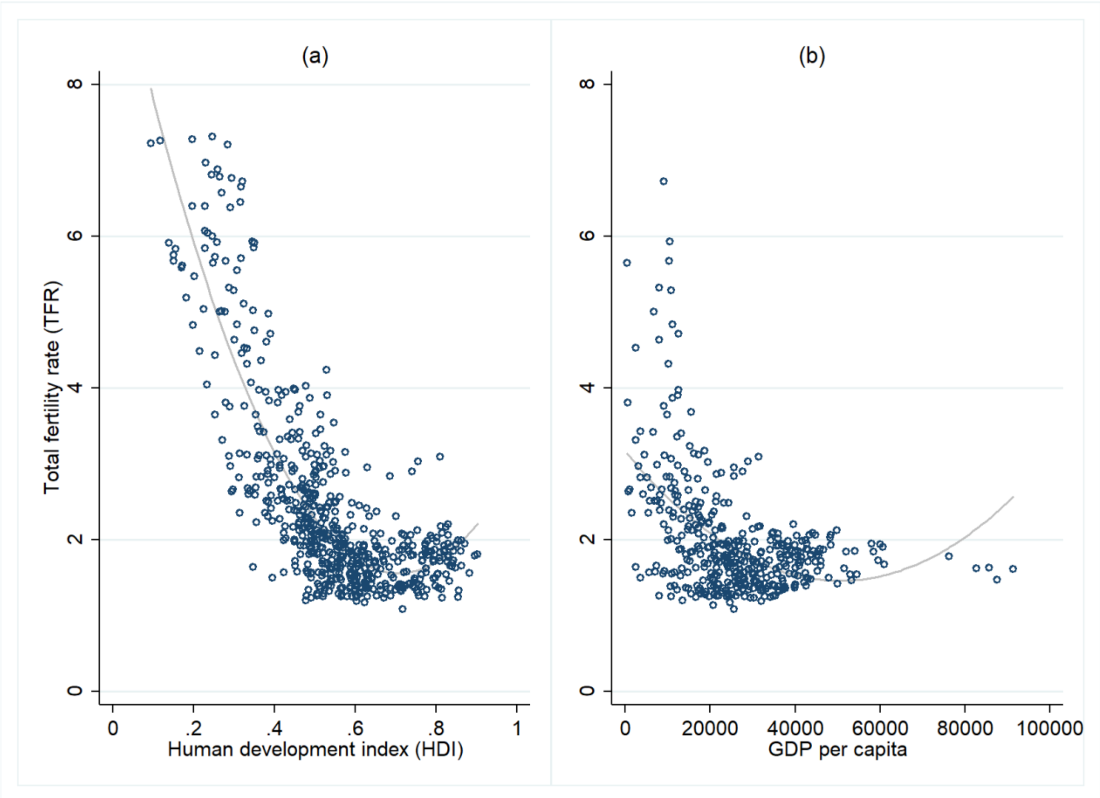

Fig. 5.1 The fertility rebound in terms of HDI and GDP per capita. Period 1970-2016. OECD countries plus Brazil, China, Colombia, India, Indonesia, Russia Federation, and South Africa.

Fig. 5.1(a) depicts the rebound with respect to Human Development Index (HDI) (Myrskylä et al., 2009) and Fig. 5.2(b) with respect to GDP per capita (Luci-Greulich and Thévenon, 2014, LGT hereafter). The former employed the Human Development Index (HDI) and were the first to indicate a threshold where the fertility starts to increase. The latter provided evidence on the rebound taking into account the economic component of HDI.

This paper aims to investigate the fertility rebound employing labour productivity, female labour force participation, and GDP per capita. We compare their contribution to the rebound by estimating the distance of their turning points for the two income country groups developed (high and low). We find that the distance is minimized for labour productivity.

5.2 Literature review

The fertility rebound took place in many of the developed countries but not all.[1] LGT endeavor to answer the latter by examining the rebound with respect to GDP per capita instead of HDI. They argue that "the use of a composite measurement of human development masks the particular contributions of each of the indicator's components (GDP per capita, life expectancy and school enrolment) and thus does not reveal why in some, especially highly developed countries, a rise in fertility comes along with increases in human development" (LGT: 188). They employ panel data analysis for OECD countries that spans from 1960 to 2007. Their findings suggest that the strong negative correlation between fertility and economic development does not hold. There is a point that turns the relationship into positive, though this point differs significantly between countries. Lacalle-Calderon et al. (2017) expand the aforementioned work. In particular, they use quantile regression analysis (Canay, 2011) to address whether an inverse J-shaped pattern exists, and

[1] For a theoretical approach, see Day (2016, 2018); Yakita (2018); Ohinata and Varvarigos (2019).

whether this pattern depends on development and fertility levels. Their results confirm both.

On the other hand, Furuoka (2013) questions the findings of Myrskylä et al. (2009) by applying threshold regression analysis on a sample of 172 countries over the period 1980-2009. Furuoka finds that the relationship between HDI and TFR becomes flatter for higher levels of HDI, but without any reversal signs. Harttgen and Vollmer (2014) show that the aforementioned relationship (HDI-TFR) does not hold for the revised HDI, neither for education, nor standard-of-living, or health components. Ryabov (2015) conducts a county level analysis for the US and provides some evidence of a negative association between human development and fertility.

There is also no consensus on the determinants of the reversal (or the pace of the decline). One strand of the literature argues that the rebound is an artifact of the fertility postponement rather than the impact of pronatalist policies (Goldstein et al., 2009; Sobotka and Lutz, 2011).[2] Others claim that the rebound is fostered in places with greater egalitarianism. Myrskylä et al. (2011) employ data for the period 1975-2008 to examine the role of gender inequality on the reversal. Their findings suggest that the rebound weakens if a country is ranked as highly developed – based on measures like health, education, and income – but low in gender equality. In the same vein, Esping-Andersen and Billari (2015) mention the importance of gender egalitarianism on their theoretical demographic framework which aims to reconcile the most recent evidence with theory. The latter attribute the increased rates of female labour force participation to policies that facilitate women to combine career and motherhood.

In this paper, we contribute to the aforementioned literature in more than one ways. First, we extend the previous literature on the rebound by taking into account the potential non-linearity of labour productivity and female labour force participation (conventional wisdom restricts to GDP per capita and HDI).

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[2] Bongaarts and Sobotka (2012: 110) question the importance of tempo effect on fertility change.

In this way, we extend Myrskylä et al. (2009) and LGT.[3] Second, to the best of our knowledge, there are no previous studies that examine the role of labour productivity as a driver for the fertility rebound. Third, we provide evidence for the rebound once we control for other covariates as well. The latter is missing from the previous empirical findings on the relevant literature. Finally, evidence is provided that labour force participation of women might not be as important for the fertility rebound as has been thought so far.

5.3 Why labour productivity

The threshold of the reversal was previously estimated for HDI and GDP per capita. It was found to vary significantly, particularly for the latter. This fact points to other factors being important for the reversal. We claim that this might be a misleading picture coming from the compositional nature of GDP per capita. Labour productivity is one of the components of the latter (average working per hour and employment ratio are the rest). We consider labour productivity as a potential driver for the rebound to further understand the phenomena. This section explicates in detail this consideration.

Labour productivity gauges GDP divided by the sum of working hours (Labour productivity = GDP(output)/sum of working hours).The higher the labour productivity, the higher the discrepancy between the output and the sum of working hours for the individuals of a country. Thus, higher labour productivity might imply not just an affluent economy, but an economy where individuals can enjoy more free time too. The latter could accommodate a fertility rebound after a certain level of labour productivity – taking into account fertility preferences.[4] In particular, at the first stages of labour productivity (low level) the case might be the one where output is low and the sum of working hours high. Data presented in Fig. 5.2 confirms this assumption. Thus, starting

[3] Notice that LGT include labour productivity and female employment ratio into their analysis. However, they do not examine for any non-linear effects on fertility besides GDP per capita.

[4] For fertility preferences, see Sobotka and Beaujouan (2014); Bhrolcháin and Beaujouan (2019).

from this point, one could possibly identify four hypothetical scenarios on the evolution (increase) of labour productivity.[5]

Scenario 1: labour productivity could increase because the output increases (sum of working hours=constant).

Scenario 2: labour productivity could increase because the sum of working hours decreases (output=constant).

Scenario 3: labour productivity could increase because output increases at a faster rate than the sum of working hours (or output decreases at a slower rate than the sum of working hours).

Scenario 4: labour productivity could increase because output increases and the sum of working hours decreases.

[5] We refer only to the long-run increases of labour productivity as being supported by the data. We do not consider short-run fluctuations. Moreover, the analysis that follows could be expanded to socio-economic differences between individuals (micro-perspective). Both are left for a future work.

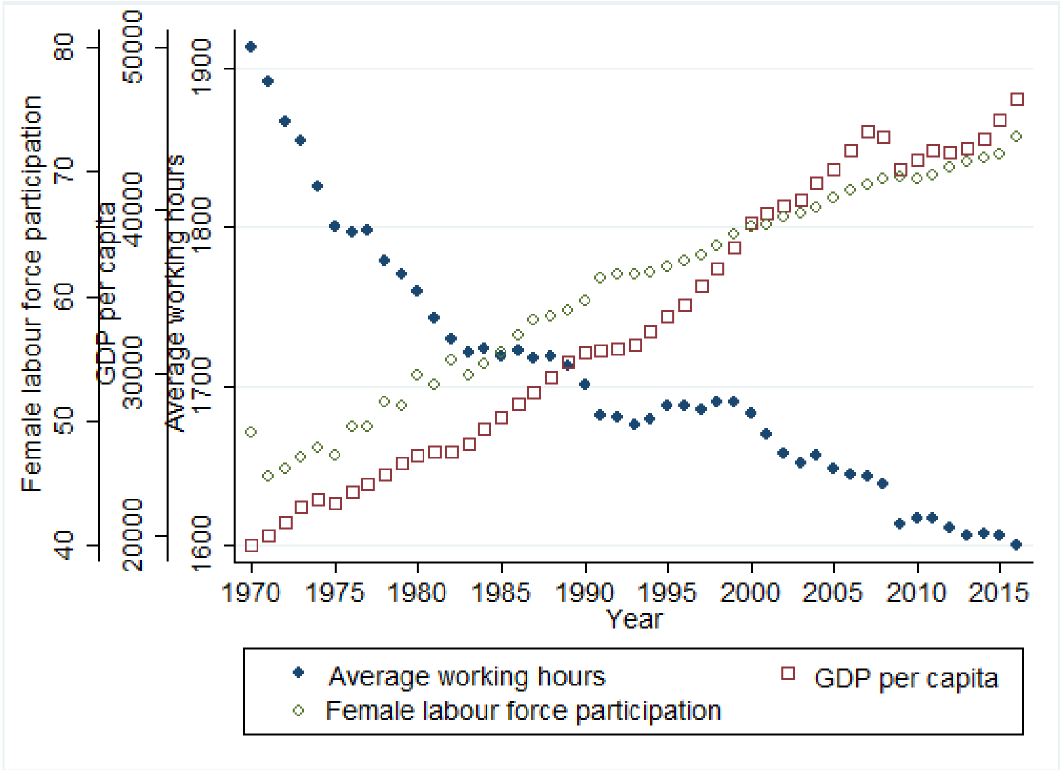

Fig. 5.2 GDP per capita, female labour force participation, and the average working hours through the years. Sample 1970-2016. Higher income countries of OECD. See Table A5.1 (Appendix A) for the countries included.

Scenario one would imply an increase in wealth without benefits in leisure time. According to Becker et al. (1990), the increasing returns to human capital will engage parents into a quality-quantity (Q-Q) trade-off and thus to substitute quantity for quality of their children. As a consequence, they would provide a higher level of education to their children by reducing their number. This would lead to a steady reduction of the fertility rate with low odds to a reversal. Moreover, SDT indicates an increase of the individualism in modern societies (self-fulfillment, marriage retreats, cohabitation and divorce increase, etc.). The latter implies that even in the absence of any Q-Q trade-off – let alone within it – individuals would prefer to reduce their fertility in order to enjoy their higher earnings. However, note that this scenario may not totally exclude the possibility of a reversal if income is high enough with respect to parents' preferences on leisure, child education, and the level of educational costs.

The increase of labour productivity for the scenario number two would probably never lead to any positive impact on fertility. According to this, individuals work less hours and thus they enjoy more free time. However, incomes remain steady (no income effect) and they are probably more hesitated to childrearing.

In scenario three, both output and working hours move towards the same direction although at different rates. Thus, in case of an increasing rate, we argue in the same way as in scenario one, whereas in case of a decreasing rate, scenario two takes place.

Scenario four could be thought as the most suitable condition for the fertility rebound. A combined increase of incomes along with the reduction of working hours would make individuals more capable to fulfill their fertility preferences, and simultaneously, enjoy their free time as we explain in more detail below.[6]

Fig. 5.2 shows the negative association between GDP per capita, women's labour participation, and the sum of the average working hours for the high income countries (henceforth HIC), where the rebound is more pronounced. What we see is an increase of labour productivity coming from both sides: GDP per capita as well as the average working hours. Hence, data propel scenario number four.

According to the analysis preceded, we think that it has become clear why the other two components of the GDP per capita (average working per hour and employment ratio) can not been seen in the same framework for the rebound. Still an important question remains: in which way would labour productivity result in an inverse J-shaped with respect to fertility? To answer, we refer to three possible stages that may explain it.

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[6] Higher earnings would mean that potential mothers are more capable to purchase childcare and thus reduce the opportunity cost they face by their increased labour force participation. However, this might not be the case for the majority of the families so that a rebound, in this case, may not occur (see also footnote 7). In contrast, a steady decline trend on working hours reduces opportunity cost without imposing new ones. For a theoretical approach on childcare purchase and fertility, see Day (2004).

Stage 1 (high fertility regime): taking into account Fig. 5.2, one might argue as follows: at the low levels of labour productivity, returns to human capital are low enough (low level of GDP per capita) so that Q-Q trade off is still absent, or very rare. In addition, female labour participation is not widely expanded in the society so that the opportunity cost of childrearing is low (rare) too. Accordingly, fertility rates remain high.

Stage 2 (steady decline regime): since GDP per capita starts to increase, returns to human capital and female labour participation are getting higher. Simultaneously, the sum of working hours decreases. The former could cause a decline in the fertility rates through the Q-Q trade off and the increased opportunity cost that women face. Moreover, the reduced working hours may accelerate the decrease – at the beginning – if we consider a parallel rise of individualism. Thus, individuals at this stage provide their children with more education and prefer to enjoy their higher earnings within an increasing free time. Fertility rates start to decline.

Stage 3 (rebound regime): as GDP per capita, female labour force participation and working hours continue their courses, we eventually reach a turning point. Beyond this point, further increases of GDP per capita could provide the capacity to parents to purchase childcare, and moreover, to fulfill their fertility preferences within the Q-Q trade-off mechanism. This would imply that they can offer the same (or higher) level of education to more children. On the other hand, the continuous decline of working hours could probably break up the competition between leisure and childrearing time (think of that in a more individualistic society as suggested by SDT). A reduction of working hours allows women to combine work, leisure, and childrearing in a more effective way.[7] Hence, less working hours might be more crucial for the rebound than

[7] Even in a developed country, the majority of families might not be able to afford childcare purchase. The percentage of these families depends on levels of inequality which according to recent evidence relates positively to growth (Bar et al., 2018). If we do not account for the reduction of working hours along with GDP per capita increments, the rebound weakens.

childcare purchase or policy reforms, which target to reconcile work and motherhood.[8] This stage implies that women's labour participation cannot be solely thought as a necessary condition for the rebound. It turns up that it depends on labour productivity levels.

Therefore, the analysis indicates a point which is crucial for the rebound and is defined from both GDP per capita and the sum of working hours. It is important to note the co-existence and co-movement of both variables. According to the three stages described, we acknowledge labour productivity and female labour force participation as the main forces that shape the rebound. We have provided some thoughts on the former due to its lack from the relevant literature. However, we do not proceed to a similar analysis for women's labour participation since it has already been suggested as one of the most significant factors on the fertility evolution (Ahn and Mira, 2002; Salamaliki et al., 2013; He and Zhu, 2016; Yakita, 2018), and particularly for the rebound: "we find a positive association between female employment and fertility for within-country variations. This implies that a change in the impact of economic development on fertility from negative to positive is only likely to happen in those countries where economic development has come along with increases in female employment." (LGT: 211).

5.4 Data and methodology

The data were retrieved from the OECD except for the variable on educational attainment that comes from the World Bank (see Appendix A, Table

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This is, perhaps, one of the reasons why we see a better fit of the rebound with respect to labour productivity rather than to GDP per capita for the low income countries (see subsection 5.6.1).

[8] Notice a possible differential effect when less working hours come from economic development, and when they arise from fertility related policies. The former would affect both parents whereas the latter mostly mothers. The investigation of this issue could be conducted in a future work.

A5.2 for all the details). GDP per capita can be decomposed into three constituents (labour productivity, average working per hour, and employment ratio) as in LGT.[9]

The sample is split into two income groups: high and low (Appendix, Table A5.1). The criterion used for the separation is the GDP per capita. We calculated the median GDP per capita for each country of the sample for the period 1970-2016, and then the average of the calculated median of all countries. The threshold found amounts to 23,480.73 in USD constant prices 2010 PPPs. Countries were divided above (HIC) and below this threshold (low income, LIC henceforth).

First, total fertility rate is regressed on GDP per capita, labour productivity, and female labour force participation for the two income country groups. Next, we estimate the mean value of the turning points for each case. The goal is to reveal the factor for which the distance of the mean turning points between low and high income countries is at a minimum, or statistically insignificant. In this context, one could claim that the factor that minimizes the distance between the mean turning points of the two groups drives the rebound.

To correct for possible bias, control variables have been included in the analysis (Appendix, Table A5.2). Hence, tertiary education accounts for human capital effects (Myrskylä et al., 2009). Life expectancy aims to capture demographic effects (Boongarts, 2017). Divorce rate has been considered to quantify cultural effects (Madsen et al., 2018). Finally, paid leave length[10] has been included to control for policy effects (Bartel et al., 2018). We should mention that our choice is related to data time – country availability. Decade dummies have been also employed following LGT. The natural logarithm transformation was used for the independent variables (Lacalle-Calderon et al., 2017).

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[9] See also Guest & Swift (2008).

[10] Length of paid maternity and parental leave available to mothers in weeks. Downloaded from https://stats.oecd.org/Index.aspx?DataSetCode=FAMILY#.

5.5 Empirical analysis

We employ panel data fixed effects (FE) and two stage least squares (2SLS) (see Schaffer, 2015 for more). Fixed effects are preferred to random effects since the analysis takes place at the country level. By the use of FE, any time-constant variable in the error term will be eliminated. To the extent that these variables are correlated to the independent ones of the model, endogeneity needs to be confronted. However, endogeneity might arise from time-varying variables being left out of the model. Therefore, in the cases examined we also control for other factors such as human capital, demographic, cultural, and policy effects. Time-varying factors that are missing from the regression and possibly affect fertility are captured by the inclusion of decade dummies. In order to further tackle potential endogeneity, we apply the 2SLS FE estimator. The instruments used are the two years lag of the linear and the quadratic term in each case.

The equation to be estimated is the following:

$$TFR_{i,t} = b_0 + b_1 lnY_{i,t} + b_2 ln\,(Y_{i,t})^2 + \sum_{j=3}^{6} b_j X'_{i,t} + \sum_{t=1}^{3} b_{7t} dummy_t + \mu_i + \varepsilon_{i,t} \qquad (1)$$

where, $TFR_{i,t}$ denotes the total fertility rate, $Y_{i,t}$ is the natural logarithm of GDP per capita, labour productivity and female labour force participation according to the case examined, $X'_{i,t}$ represents the control variables: tertiary educational attainment, life expectancy, divorce rate, parental paid leave length, $dummy_t$ accounts for the three decade dummies (t=1,2,3) and takes the value 1 ($dummy_t = 1$) for the decade of reference (otherwise takes the value 0), μ_i is the fixed-effect term and $\varepsilon_{i,t}$ is the stochastic error term. The turning points have been calculated as $-\frac{b_1}{2b_2}$, where b_1 is the coefficient of $lnY_{i,t}$ and b_2 of $ln\,(Y_{i,t})^2$.[11]

[11] The turning point occurs to the point where the first derivative equals to zero: $TFR = b_1 lnGDP + b_2 \ln(GDP)^2 \Rightarrow \frac{d(TFR)}{dln\,(GDP)} = 0 \Rightarrow b_1 + 2b_2 ln(GDP) = 0 \Rightarrow ln(GDP) = -\frac{b_1}{2b_2}$.

5.6 Results

5.6.1 Preliminary analysis

We plot the relationship between TFR and GDP per capita (Fig. 5.3), labour productivity (Fig. 5.4), and female labour force participation (Fig. 5.5) for LIC and HIC, to examine the factors that are characterized by a reversal and compare between the two groups.

Fig. 5.3 presents the reversal of TFR as GDP per capita increases for the HIC group. TFR starts slightly to rise from 1.6 after the level of 30,000 to 1.8 at the 45,000 (USD constant prices 2010 PPPs). The LIC group experiences a (weaker) rebound but at a lower level of GDP per capita than that of HIC. These results are partly – due to sample differences – aligned to Bongaarts (2008) who find some evidence on the rebound for developing countries too. However, they are opposed to findings by Furuoka (2013) and Harttgen and Vollmer (2014).

In Fig. 5.4 (labour productivity), we observe that both HIC and LIC are similarly shaped. In contrast to Fig. 5.3 (GDP per capita), the rebound occurs approximately at the same (or closer) levels of labour productivity. Lastly, in Fig. 5.5 (female labour force participation), the rebound is present only for HIC.

Overall, visual inspection allows us to make the following points: (1) the rebound is more distinct for HIC when female labour force participation is considered and (2) the rebound is similar to both groups for labour productivity.

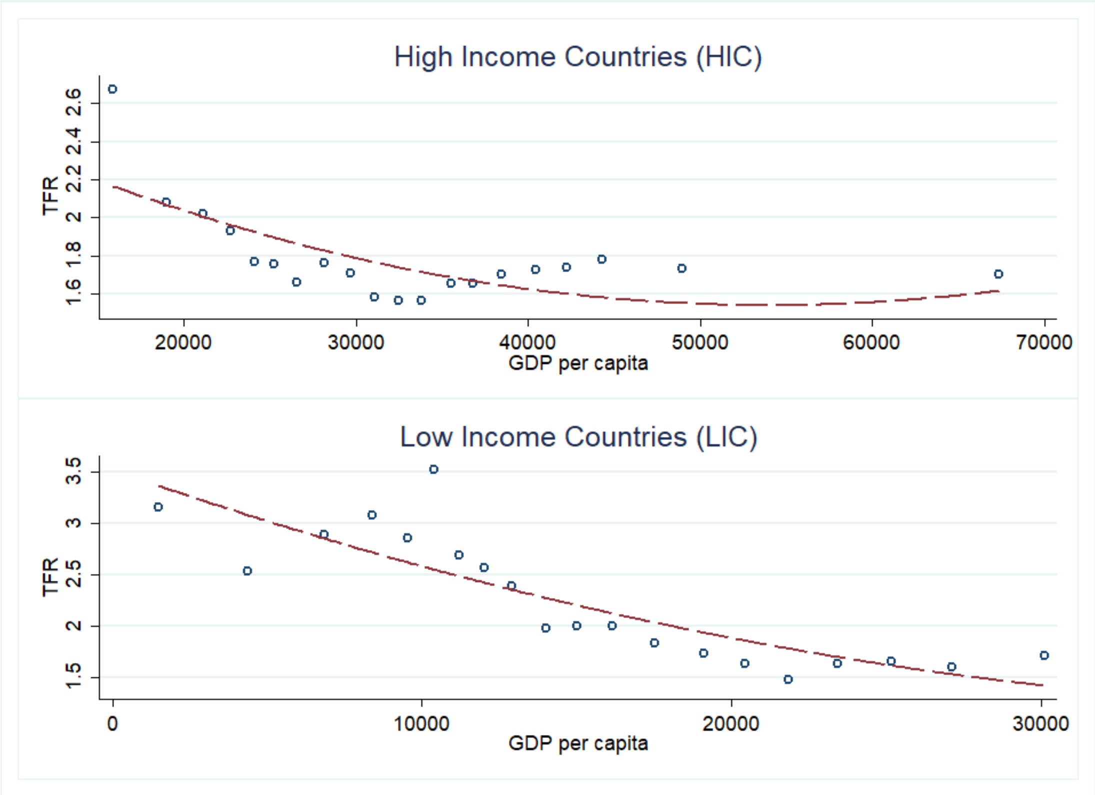

Fig. 5.3 Panel plots between TFR and GDP per capita.

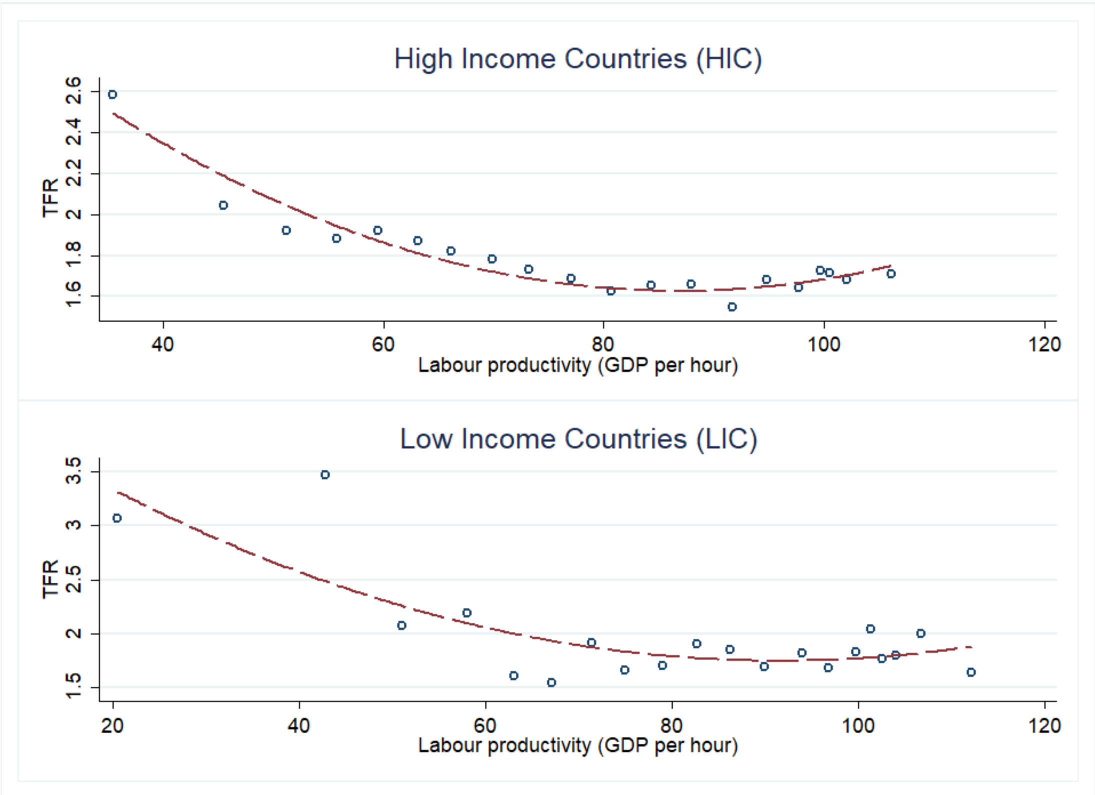

Fig. 5.4 Panel plots between TFR and labour productivity.

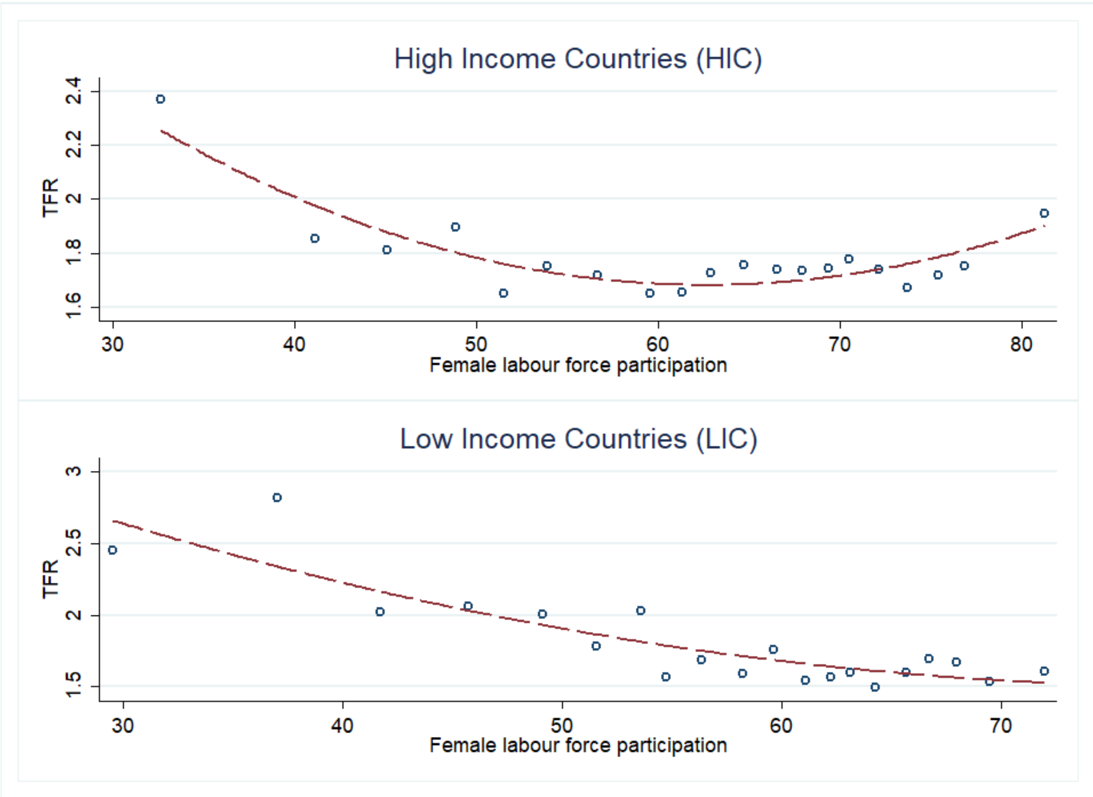

Fig. 5.5 Panel plots between TFR and female labour force participation.

5.6.2 Regression results

This subsection uses regression analysis to examine the presence of the fertility rebound with respect to GDP per capita, labour productivity, and female labour force participation.[12] We consider three cases: (A) no decade dummies or other control variables, (B) only decade dummies, and (C) decade dummies and control variables. Next, we estimate the turning points of the observed rebound and obtain the difference (distance) of the turning points between the two income groups for each factor. Lastly, we test for the statistical significance of the estimated differences.

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[12] We also considered HDI but the short sample that is available (1990 onwards) does not provide a lengthy period to examine the rebound and the turning points (http://hdr.undp.org/en/data). We have also considered the HHDI (historical HDI – https://espacioinvestiga.org/home-hihd/hihd-downloads/?lang=en) which is a 5-year index. Results on the latter are similar to those obtained for labour productivity as far as the distance is concerned. They are available upon request.

Table 5.1 summarizes the regression results[13] obtained from equation 1 for both HIC and LIC groups. With respect to GDP per capita, rebound is confirmed at the 1% level of significance for the HIC group independent of the case examined (A,B,C) or the method applied (FE, 2SLS FE). On the LIC, the rebound is present and statistically significant at the 1% level; an exception comes from case B (10%) and C (5%) in FE. On labour productivity results (GDP per hour), rebound is present for both groups. When FE is applied, it becomes statistically significant at the 1% in cases A and B for the HIC, and in A for the LIC. Controlling for endogeneity, labour productivity becomes statistically significant at the 1% level in all cases for both income groups. Finally, the rebound with respect to the female labour force participation is present at the 1% level of significance for the HIC group apart from case C of the FE where no convexity found. The LIC group is not characterized by a convex relationship with respect to female labour participation other than case B of FE, at the 5% level of significance.

Table 5.1 Summary of the regression results.

		FE	
Case	GDP per capita	GDP per hour	Flfp
A	HIC*** and LIC***	HIC*** and LIC***	HIC***
B	HIC*** and LIC*	HIC*** and LIC	HIC*** and LIC**
C	HIC*** and LIC**	HIC and LIC	linear
		2SLS FE	
Case	GDP per capita	GDP per hour	Flfp
A	HIC*** and LIC***	HIC*** and LIC***	HIC***
B	HIC*** and LIC***	HIC*** and LIC***	HIC***
C	HIC*** and LIC***	HIC*** and LIC***	HIC***

Note: A: no decade dummies or control variables. B: decade dummies. C: decade dummies and control variables. *Flfp* stands for female labour force participation. The presence of HIC or LIC in the Table implies a convex relationship (rebound) for the corresponding group. Asterisks represent the post estimation *F*-test conducted for the linear and the quadratic term. *** *p-value*< 0.01, ** *p-value*< 0.05, * *p-value*< 0.1.

[13] Detailed regression results are shown in Appendix B.

Table C5.1 (Appendix) presents the calculated turning points in the cases where the rebound confirmed (independent of its statistical significance). We focus on their differences and present them in Table 5.2. We observe that the difference between the mean turning points of the two income groups is minimized for labour productivity: 0.109<2.015 and 0.810, 0.552<1.068, 0.747<11.860. Second comes the labour force participation of women and last the GDP per capita (0.109<0.810<2.015). On the derived statistical tests for the FE, there are two cases where the null hypothesis of no difference between the two mean turning points cannot be rejected. The one concerns the GDP per capita (A) and the other comes from labour productivity [GDP per hour (B)].

Table 5.2 Differences in turning points (in natural logarithms[14]).

	FE			2SLS FE		
	(1)	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**
Case	GDPpc	GDPph	Flfp	GDPpc	GDPph	Flfp
A	11.860	0.747***	–	67.470	1.035	–
B	2.015***	0.109	0.810***	2.401	0.248	–
C	1.068***	0.552**	–	1.148	0.898	–

Note: A: no decade dummies or control variables, B: decade dummies, C: decade dummies and control variables. *Flfp* stands for female labour force participation, *GDPpc* for GDP per capita, and *GDPph* for GDP per hour. Columns 1 - 3 refer to FE and 4 - 6 to 2SLS FE. Null hypothesis: no difference between the mean turning points. *** *p-value*< 0.01, ** *p-value*< 0.05, * *p-value*< 0.1.

However, note that the turning point for LIC of the former had not been found statistically significant (Table C5.1, Appendix). The opposite holds for labour productivity. This implies that the distance for the latter is at a minimum among all factors examined, and moreover, the statistical significance test

[14] The natural logarithmic transformation does not change the results that would be obtained for the differences in levels. Logarithm is a monotonically increasing function and turning points are positive, thus: $x_1 > x_2 \Leftrightarrow lnx_1 > lnx_2(1)$, $\beta - \alpha > \gamma - \delta \Leftrightarrow \beta > \alpha$ and $\gamma > \delta(2)$, thus, $\ln(\beta) - \ln(\alpha) > \ln(\gamma) - \ln(\delta) \Leftrightarrow \ln\left(\frac{\beta}{\alpha}\right) > \ln\left(\frac{\gamma}{\delta}\right) \Leftrightarrow (1) \Leftrightarrow \frac{\beta}{\alpha} > \frac{\gamma}{\delta}$, which is valid due to (2).

indicates that there is no difference. The statistical (in)significance for the differences of the 2SLS estimator is depicted in Fig. 5.6-5.8 (Appendix D) where we plot the confidence intervals. Again, the distance for labour productivity is statistically insignificant in cases B (Fig. 5.7) and C (Fig. 5.8) since the confidence intervals overlap. The null hypothesis could not be rejected only in case A although the difference still remains the lowest for labour productivity. On the other hand, confidence intervals do not overlap for GDP per capita and hence the null of no difference can be rejected. This picture changes only in case A (Fig. 5.6) where GDP per capita overlaps – in line with the results on FE of Table 5.2.

5.6.3 Robustness check

We consider two robustness checks. First, FE with Discroll-Kraay standard errors which aim to control for general forms of cross-sectional dependence (Driscoll and Kraay, 1998), and second, the GMM difference estimator. Arellano and Bond (1991) estimator (one step[15] difference GMM) removes the fixed country-specific effects by taking first differences. Moreover, we deal with the inflation of the gaps created by the first difference transformation (Roodman, 2009: 21). A control for serial correlation and heteroskedasticity has been imposed. Robustness results confirm to a great extent the ones preceded (see Appendix C, Tables C5.2-C5.5). Detailed regression results and confidence intervals plots are available upon request.

5.7 Conclusions

This paper tries to shed more light on the fertility rebound. We test for the presence of the rebound taking into account GDP per capita, labour productivity, and female labour force participation. OECD countries are split into two income groups (high and low). We focus on the turning points of the reversal for each factor and estimate their distance between the two groups

[15] For a discussion on one and two step GMM estimators, see Hwang and Sun (2018).

considered. Labour productivity found to min the distance, and hence, it could be thought as the most important factor for the onset of the rebound.

Results reveal that the fertility rebound is present for both LIC and HIC in terms of labour productivity. Rebound also holds (mainly) in terms of female labour force participation for the HIC group. Most important, difference in the turning points of the rebound is closer between the two income country groups when labour productivity is considered. The latter implies that the rebound occurs when countries reach a specific range of values of labour productivity rather than GDP per capita or female labour force participation. Therefore, we could infer that labour productivity arises as the most important economic driver for the onset of the fertility rebound. Female labour force participation follows. Finally, labour productivity might provide us an explication on why fertility rebound has been also observed in countries with low levels of female labour force participation.

5.8 References

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5.9 Appendix

A. Descriptive statistics

Table A5.1 Country separation in high and low income groups.

Country (LIC)	GDP per capita (mean)	Country (HIC)	GDP per capita (mean)
China	2,036.33	NewZealand	23,572.12
India	3,474.685	Spain	23,782.09
Indonesia	6,568.439	Slovenia	24,097.82
Colombia	8,557.774	United Kingdom	26,222.4
South Africa	10,444.47	Finland	26,853.29
Turkey	11,165.9	Iceland	28,594.84
Brazil	12,500.98	France	29,377.84
Mexico	12,982.01	Sweden	30,019.06
Chile	13,617.6	Australlia	30,211.21
Korea	14,137.66	Belgium	30,675.57
Poland	15,280.41	Italy	30,700.02
Latvia	16,873.54	Japan	30,968.96
Lithuania	18,490.18	Canada	31,243.55
SlovakRepublic	18,593.09	Austria	31,447.6
RussianFed.	19,250.53	Germany	32,089.46
Hungary	20,148.36	Netherlands	32,783.76
Estonia	20,350.21	Denmark	33,968.68
Greece	20,596.22	United States	36,746.48
Israel	21,127.55	Norway	43,033.44
Portugal	21,157.29	Switzerland	43,886.07
Ireland	22,456.31	Luxembourg	57,034.16
CzechRepublic	22,553.57		

Note: Countries are presented from the lowest GDP per capita mean (China) to the highest (Luxembourg). GDPper capita measured in USD constant prices 2010 PPPs. *LIC* stands for low income countries. *HIC* stands for high income countries.

Table A5.2 Variables employed in the analysis.

	Source	Time frame
Total fertility rate	OECD	1970–2016
GDP per capita	OECD	1970–2016
GDP per hour	OECD	1970–2016
Female labour force participation	OECD	1970–2016
Tertiary education (both genders)	World Bank	1970–2016
Life expectancy	OECD	1970–2016
Divorce rate	OECD	1970–2016
Length of paid maternity and parental leave	OECD	1970–2016

Note: Time starts at different years for each country (unbalanced panel).

B. Regression results

Table B5.1 FE regression results on cases A and B.

	High OECD Income Countries					
Dependent: TFR	(1)	(2)	(3)	(4)	(5)	(6)
ln(GDP per capita)	-19.82***			-17.42***		
	(5.410)			(4.304)		
ln(GDP per capita)^2	0.934***			0.810***		
	(0.262)			(0.208)		
ln(labour productivity)		-8.818***			-7.943***	
		(1.532)			(2.448)	
ln(labour productivity)^2		0.988***			0.897***	
		(0.183)			(0.298)	
ln(Flfp)			-13.33***			-13.15***
			(4.508)			(4.420)
ln(Flfp)^2			1.562**			1.582***
			(0.552)			(0.541)
Decade dummies	No	No	No	Yes	Yes	Yes
Constant	106.7***	21.36***	30.10***	95.22***	19.38***	29.26***
	(27.93)	(3.214)	(9.206)	(22.28)	(5.001)	(9.089)
R-squared	0.520	0.470	0.407	0.560	0.491	0.503
Observations	986	960	866	986	960	866
Number of countries	21	21	21	21	21	21

Note: For Tables B5.1 – B5.6: *** *p–value*< 0.01, ** *p–value*< 0.05, * *p–value*< 0.1. Instrumental variables: two years lag of the linear and the quadratic term. *Flfp* stands for female labour force participation.

Table B5.2 FE regression results on cases A and B.

	Low OECD Income Countries					
Dependent: TFR	**(1)**	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**
ln(GDP per capita)	-1.580*			-2.578**		
	(0.904)			(1.023)		
ln(GDP per capita)^2	0.0352			0.148**		
	(0.0570)			(0.0606)		
ln(labour productivity)		-4.560***			-2.337	
		(1.283)			(1.935)	
ln(labour productivity)^2		0.438**			0.258	
		(0.194)			(0.258)	
ln(Flfp)			9.601			-5.272
			(8.482)			(4.870)
ln(Flfp)^2			-1.332			0.787
			(1.088)			(0.603)
Decade dummies	No	No	No	Yes	Yes	Yes
Constant	14.02***	13.42***	-15.21	14.98***	8.628**	11.53
	(3.567)	(2.039)	(16.54)	(4.377)	(3.245)	(9.864)
R-squared	0.342	0.546	0.097	0.728	0.754	0.480
Observations	729	460	478	729	460	478
Number of countries	22	16	21	22	16	21

Note: The number of countries varies due to data availability.

Table B5.3 FE regression results on case C.

	High OECD Income Countries			Low OECD Income Countries		
Dependent: TFR	(1)	(2)	(3)	(4)	(5)	(6)
ln(GDP per capita)	-12.01***			-13.97**		
	(3.052)			(5.231)		
ln(GDP per capita)^2	0.561***			0.725**		
	(0.150)			(0.268)		
ln(labour productivity)		-4.720			-2.855	
		(3.235)			(2.817)	
ln(labour productivity)^2		0.510			0.351	
		(0.383)			(0.327)	
ln(Flfp)			-11.15***			0.103
			(3.813)			(2.318)
ln(Flfp)^2			1.426***			0.0253
			(0.481)			(0.320)
ln(education)	0.0159	0.175	0.0447	-0.220***	-0.121	-0.0976
	(0.155)	(0.128)	(0.154)	(0.0569)	(0.103)	(0.0933)
ln(lifeexpectancy)	-1.034	-2.128	-3.261	-3.394**	-2.248	-1.627
	(1.665)	(2.421)	(1.974)	(1.119)	(1.805)	(1.760)
ln(divorce rate)	-0.252***	-0.232**	-0.274**	0.0388	-0.0725	-0.109
	(0.0820)	(0.0902)	(0.115)	(0.0306)	(0.0587)	(0.0674)
ln(paid leavelength)	0.0686	0.113**	0.0118	-0.131**	-0.0859	-0.105**
	(0.0675)	(0.0506)	(0.0503)	(0.0433)	(0.0583)	(0.0388)
Decade dummies	Yes	Yes	Yes	Yes	Yes	Yes
Constant	70.16***	20.93**	37.50***	85.14***	18.22*	9.007
	(17.50)	(7.523)	(11.20)	(26.62)	(9.353)	(9.891)
R-squared	0.463	0.365	0.488	0.861	0.811	0.753
Observations	581	557	503	236	215	214
Number of countries	19	19	19	11	10	11

Note: Columns 1 – 3 refer to HIC and columns 4 – 6 refer to LIC. The number of countries varies due to data availability.

Table B5.4 2SLS FE regression results on cases A and B.

	High OECD Income Countries					
Dependent: TFR	**(1)**	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**
ln(GDP per capita)	-19.35***			-17.37***		
	(1.585)			(1.414)		
ln(GDP per capita)^2	0.916***			0.811***		
	(0.0761)			(0.0672)		
ln(labour productivity)		-9.771***			-9.408***	
		(0.604)			(0.886)	
ln(labour productivity)^2		1.105***			1.074***	
		(0.0707)			(0.110)	
ln(Flfp)			-12.26***			-13.40***
			(1.740)			(1.750)
ln(Flfp)^2			1.454***			1.645***
			(0.216)			(0.221)
Decade dummies	No	No	No	Yes	Yes	Yes
R-squared	0.447	0.406	0.333	0.485	0.421	0.423
Observations	944	918	821	944	918	821
Number of countries	21	21	21	21	21	21

Table B5.5 2SLS FE regression results on cases A and B.

	Low OECD Income Countries					
Dependent: TFR	**(1)**	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**
ln(GDP per capita)	-0.964*			-2.455***		
	(0.525)			(0.486)		
ln(GDP per capita)^2	0.00617			0.148***		
	(0.0291)			(0.0283)		
ln(labour productivity)		-3.795***			-1.623**	
		(0.512)			(0.734)	
ln(labour productivity)^2		0.348***			0.175*	
		(0.0673)			(0.0951)	
ln(Flfp)			12.24***			1.004
			(4.138)			(3.255)
ln(Flfp)^2			-1.645***			-0.0512
			(0.521)			(0.411)
Decade dummies	No	No	No	Yes	Yes	Yes
R-squared	0.301	0.498	0.080	0.719	0.728	0.363
Observations	685	428	429	685	428	429
Number of countries	22	16	21	22	16	20

Note: The number of countries varies due to data availability.

Table B5.6 2SLS FE regression results on case C.

	High OECD Income Countries			Low OECD Income Countries		
Dependent: TFR	(1)	(2)	(3)	(4)	(5)	(6)
ln(GDP per capita)	-11.64***			-15.90***		
	(1.724)			(3.339)		
ln(GDP per capita)^2	0.549***			0.840***		
	(0.0802)			(0.172)		
ln(labour productivity)		-5.659***			-2.664	
		(1.434)			(1.620)	
ln(labour productivity)^2		0.618***			0.362*	
		(0.174)			(0.189)	
ln(Flfp)			-12.76***			1.346
			(1.907)			(2.114)
ln(Flfp)^2			1.647***			-0.151
			(0.242)			(0.278)
ln(education)	0.00471	0.180***	0.0772*	-0.220***	-0.00515	0.0422
	(0.0552)	(0.0435)	(0.0459)	(0.0735)	(0.0555)	(0.0676)
ln(lifeexpextancy)	-1.166	-1.820	-3.707***	-4.776***	-4.403***	-1.842**
	(0.963)	(1.226)	(0.861)	(1.171)	(1.134)	(0.831)
ln(divorce rate)	-0.238***	-0.221***	-0.251***	0.0298	-0.0790**	-0.177***
	(0.0488)	(0.0613)	(0.0474)	(0.0334)	(0.0365)	(0.0555)
ln(paid leavelength)	0.0635***	0.122***	0.00637	-0.155***	-0.0926***	-0.0942***
	(0.0218)	(0.0239)	(0.0204)	(0.0311)	(0.0269)	(0.0228)
Decade dummies	Yes	Yes	Yes	Yes	Yes	Yes
R-squared	0.407	0.300	0.470	0.847	0.823	0.733
Observations	570	545	486	225	205	203
Number of countries	19	19	19	10	10	10

Note: Columns 1 – 3 refer to HIC and columns 4 – 6 refer to LIC. Number of countries varies due to data availability.

C. Turning points and robustness check

Table C5.1 Turning points (in natural logarithms).

	Low Income Countries					
	FE			**2SLS FE**		
	(1)	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**
Case	GDP per capita	GDP per hour	Flfp	GDP per capita	GDP per hour	Flfp
A	22.465	5.210***	concave	78.039	5.455***	concave
B	8.736***	4.537***	3.347***	8.312***	4.626***	concave
C	9.631***	4.072***	linear	9.463***	3.6780***	concave
	High Income Countries					
	FE			**2SLS FE**		
	(1)	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**
Case	GDP per capita	GDP per hour	Flfp	GDP per capita	GDP per hour	Flfp
A	10.605***	4.463***	4.268***	10.569***	4.420***	4.214***
B	10.752***	4.428***	4.158***	10.714***	4.378***	4.071***
C	10.699***	4.624***	3.908***	10.6112***	4.576***	3.874***

Note: A: no decade dummies or control variables, B: decade dummies, C: decade dummies and control variables. Columns 1 - 3 refer to FE and 4 - 6 to 2SLS FE. *** *p-value*< 0.01, ** *p-value*< 0.05, * *p-value*< 0.1. *Flfp* stands for female labour force participation.

Table C5.2 Robustness check – turning points (in natural logarithms).

	Low Income Countries					
	CSD control			**GMM difference**		
	(1)	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**
Case	GDP per capita	GDP per hour	Flfp	GDP per capita	GDP per hour	Flfp
A	22.465	8.736***	9.631***	3.668	.0667	2.801*
B	5.210***	4.537***	4.072 ***	3.215	concave	concave
C	concave	3.347***	linear	9.388***	4.005***	4.004***
	High Income Countries					
	CSD control			**GMM difference**		
	(1)	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**
Case	GDP per capita	GDP per hour	Flfp	GDP per capita	GDP per hour	Flfp
A	10.605***	10.752***	10.700***	9.554***	4.196***	3.526***
B	4.463***	4.428***	4.624***	8.485***	4.049***	3.396***
C	4.268***	4.157***	3.908***	concave	4.026***	3.663***

Note: A: no decade dummies or control variables, B: decade dummies, C: decade dummies and control variables. Columns 1 - 3 refer to CSD and 4 - 6 to GMM. *** *p-value*< 0.01, ** *p-value*<0.05, * *p-value*< 0.1. *Flfp* stands for female labour force participation.

Table C5.3 Robustness check – summarized regression results.

CSD control			
Case	GDP per capita	GDP per hour	Flfp
A	HIC*** and LIC***	HIC*** and LIC***	HIC***
B	HIC*** and LIC***	HIC*** and LIC***	HIC*** and LIC***
C	HIC*** and LIC***	HIC***	HIC***
GMM difference			
Case	GDP per capita	GDP per hour	Flfp
A	HIC** and LIC***	HIC*** and LIC***	HIC*** and LIC*
B	HIC** and LIC***	HIC***	HIC***
C	LIC***	HIC and LIC***	HIC*** and LIC

Note: A: no decade dummies or control variables, B: decade dummies, C: decade dummies and control variables. The presence of HIC or LIC in the Table implies the presence of a convex relationship (rebound) for the corresponding group. Asterisks represent the post estimation F-test conducted for both the linear and the quadratic term. *** *p-value*< 0.01, ** *p-value*< 0.05, * *p-value*< 0.1. *Flfp* stands for female labour force participation.

Table C5.4 Robustness check – differences in turning points (natural logarithms).

	CSD control			GMM difference		
	(1)	(2)	(3)	(4)	(5)	(6)
Case	GDPpc	GDPph	Flfp	GDPpc	GDPph	Flfp
A	11.860	0.747	–	5.886	4.130	0.726
B	2.015	0.109	0.810	5.270	–	–
C	1.068	0.552	–	–	0.021	0.340

Note: A: no decade dummies or control variables, B: decade dummies, C: decade dummies and control variables. *Flfp* stands for female labour force participation, *GDPpc* for GDP per capita, and *GDPph* for GDP per hour. Columns 1 - 3 refer to CSD and 4 - 6 to GMM. Null hypothesis: no difference between the mean turning points. *** *p-value*< 0.01, ** *p-value*< 0.05, * *p-value*< 0.1.

Table C5.5 Statistical significance of the values in Table C5.4 according to the confidence intervals plots.

	CSD control			GMM difference		
	(1)	(2)	(3)	(4)	(5)	(6)
Case	GDPpc	GDPph	Flfp	GDPpc	GDPph	Flfp
A	n.r.	n.r.	–	n.r.	n.r.	n.r.
B	r.	n.r.	r.	n.r.	–	–
C	r.	n.r.	–	–	n.r.	n.r.

Note: A: no decade dummies or control variables, B: decade dummies, C: decade dummies and control variables. *Flfp* stands for female labour force participation, *GDPpc* for GDP per capita, and *GDPph* for GDP per hour. Null hypothesis: no difference between the mean turning points. n.r.: null not rejected. r: null rejected. Columns 1 - 3 refer to CSD and 4 - 6 to GMM.

D. Confidence intervals for 2SLS FE

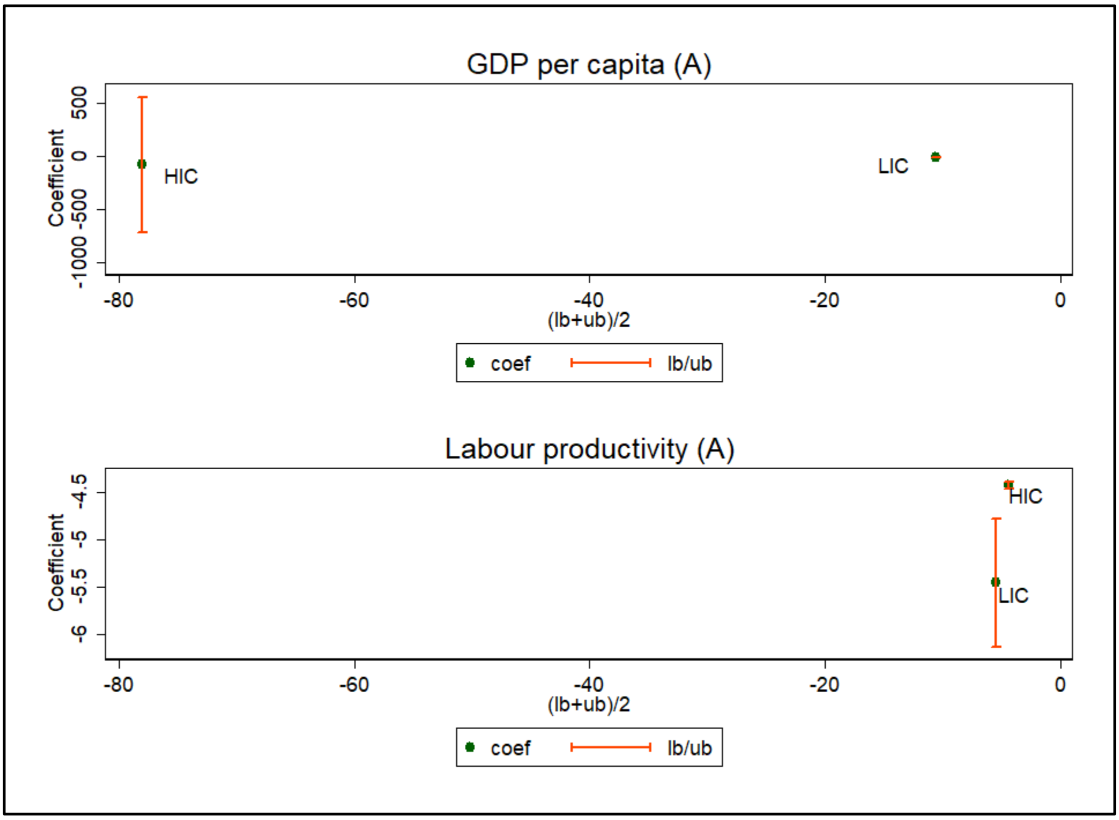

Fig. 5.6 Confidence intervals for case A.

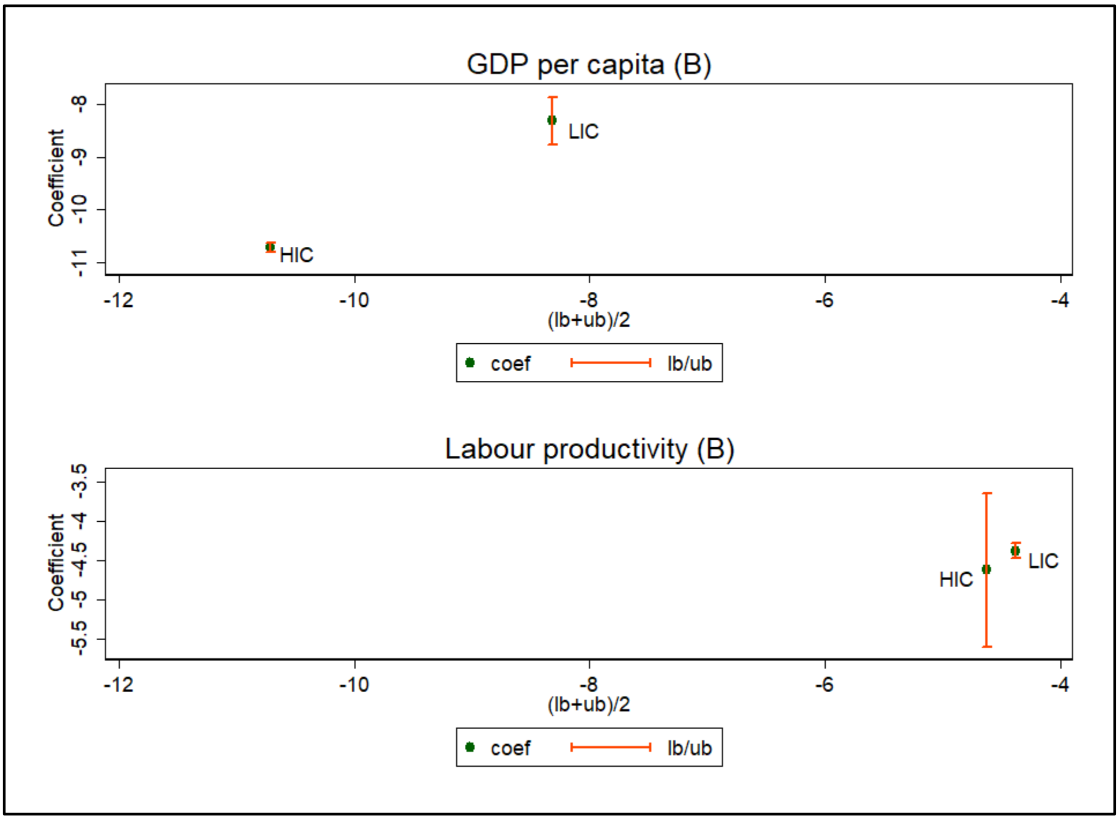

Fig. 5.7 Confidence intervals for case B.

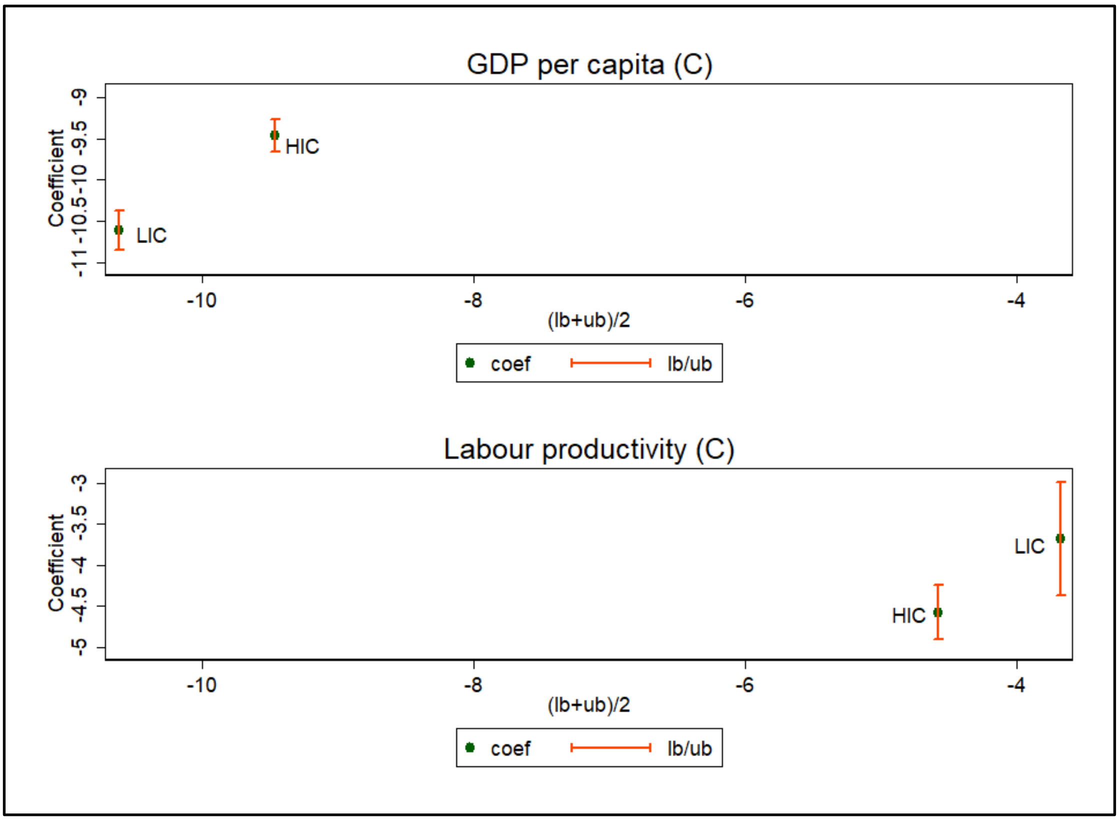

Fig. 5.8 Confidence intervals for case C.

CHAPTER 6

Epilogue

As an epilogue, we would like to make some suggestions for future research regarding the topic of each one chapter. There are still many issues left open in this thesis. What follows is not just assumptions or speculations. Many of the suggestions presented below are already under the scope, and results have been obtained in many cases.

On the second chapter

To begin with, we argue that relative income hypothesis consists of two dimensions. The first one refers to within self socio-economic comparisons (childhood-young adulthood), and the second one on social-comparisons that take place between individuals (in both childhood and young adulthood). Thus, with respect to the second chapter, where we study the issue of marriage in young adulthood, one could conduct a micro-approach paper to test the potential effect of both dimensions on the age at marriage. These two could not be separated within an aggregate analysis. Hence, one could estimate the mean/median family income in childhood for specific years (i.e. 5-15 years old) by each individual. This would be the "within comparisons" effect.

However, the "within dimension", which decisively contributes to the formation of material aspirations, is also affected by the relative socio-economic position of child's family with respect to his peers ("between dimension"). That is, feelings of relative deprivation (Dennison and Swisher, 2019) in childhood might be stronger or weaker, depending on the socio-economic position of the child. This is an important issue that the relevant literature has not mentioned regarding the Easterlin hypothesis. A stronger relative deprivation in childhood, for instance, would increase material aspirations and subsequently the threshold to be reached in young adulthood (relative income would reduce). State it differently; in case of a stronger relative deprivation effect in childhood, aspirations being represented *only* by personal childhood family income would under-estimate their true impact on the age at marriage (or any other demographic event). Aggregate analysis – as it has computed in this thesis – mixes and presents the total effect of both (personal childhood family income and relative deprivation). Hence, one could perceive it as a merit (see p. 26).

Only in a micro-analysis framework, however, one could distinguish between these two effects. The query is how to account for relative deprivation in childhood. We think that this could be done by first, separating individuals into socio-economic groups in childhood (according to the parental income and/or education) and second, by creating an aggregate measure of inequality (in county or state level). Thus, the core assumption would be the one where a higher inequality would result in greater feelings of relative deprivation experienced and transmitted to young adulthood by children of the lower socio-economic status. This would be the "between comparisons" effect.

On the third chapter

The importance of the third chapter comes from the fact that the effect of relative income found significant on the rate of children born out-of-wedlock after we controlled for marriage. Thus, it might open further insights on the relationship between relative income and non-marital fertility.

In particular, as we argue in the corresponding paragraph (p.58), cohabitation is evidenced to be more a case of individuals who belong to lower socio-economic statuses. Theoretical literature implies that non-marital fertility will be prevalence in individuals lying at the lower distributions of income (Willis, 1999), and evidence also suggest that individuals of higher educational attainment have lower propensity on premarital births (Perelli-Harris et al., 2010). Taking into account the latter and the socio-economic perspective of relative income, one could infer on the potential of socio-economic mobility of the out-of-wedlock childbearing.

Accordingly, one could proceed to a micro-level empirical analysis, and develop the relative income for the lower socio-economic status individuals (see also comments in p.12-13). Then, it could be tested the assumption of a negative correlation based on the potential of a socio-economic mobility shift: a higher relative income would decrease odds to non-marital births. Such an outcome would imply that young men who find that face more chances to climb up the social ladder would aim to avoid non-marital birth. This would be a potential link between socio-economic mobility and premarital births.

On the fourth chapter

The fourth chapter probes the fertility postponement and the increased female educational attainment as well. Here, one could further distinguish between married and single women's postponement, and test the hypothesis stated for each category separately.[1]

The latter would induce some differences. Young married women would not be affected by the relative income of young unmarried men rather than their husbands. In this context, one could first test for the relative income effect of young single men on single women's education and their corresponding fertility postponement. Next, it could be examined the case of young married men's relative income effect on the female labour force participation of young married women (instead of education), the length from marriage to first birth, and the length between subsequent births (instead of mean age at first birth). It might be more possible for a married woman, especially for a young one, to contribute to family aspirations accomplishment by working rather than attending higher education.

In the first chapter, we aimed to demonstrate that marriage fall in earlier ages could be also attributed to the decline of young men's relative income. Moreover, Fig. 2.5 in p.34 depicts the relative income differences between

[1] The author is grateful to Prof. Thanasis Stengos for useful comments received at the 5th AMEF conference.

married and single young men. Relative income found higher for the former. This evidence reveals that separating between married and single women and men, one might not expect relative income to induce the same impact in both cases. With respect to singles, we think that it would be even greater than the one presented in chapter 3, but results would be less favourable for married young adults.

On the fifth chapter

The last paper of this thesis provides some evidence on the correlation between labour productivity and the fertility rebound. The rebound has been mostly linked to countries where levels of female labour force participation are high (Luci-Greulich and Thévenon, 2014). The latter occurs in developed countries. Literature, however, does not explicate why we also observe the rebound in some developing countries (Bongaarts, 2008), where the percentage of women in labour market is still low (and subsequently egalitarianism too). Labour productivity might be an answer. In this perspective, labour productivity could be thought as the driver for the onset of the rebound. By the latter, we imply that fertility could start recovering after decades of fall because individuals enjoy more free time and higher earnings at the same time. If this combination is not in place, female labour force participation might not lead to the rebound. However, the latter is a strong assumption and difficult to be confirmed; labour productivity and women's participation in the labour market are highly correlated. Spotting the rebound with respect to labour productivity for high and low income countries, and the female labour force participation mostly for the higher income countries of the OECD, one could infer that the role of the former is to touch off the rebound, whilst the latter to found it. This is our main conclusion that might merit more investigation in future research.

As a first step, we could consider a regression quantile approach. This would give us the opportunity to observe how these two variables behave in low and high income countries across different fertility levels.

A second step would be a micro-data approach. As it is explained in detail in chapter 4, labour productivity has been tested for the rebound not just because of its origin from GDP, but because it represents – in a macro perspective – two important factors on fertility decision: time and income. This hypothesis could be further investigated at the couple (individual) level. One could estimate the ratio of the total income to the total amount of working time for each couple, and test it for the number or timing of births. Moreover, a separation regarding the socio-economic and working time regime would be even more beneficial on testing this hypothesis (e.g., a couple with high earnings and high working time – high socio-economic status – might be able to purchase childcare and substitute for working time).

Conclusion

In conclusion, this thesis investigates four demographic events: marriage, premarital births, fertility postponement (with female education), and the fertility rebound. The cohesive element among the four chapters of this thesis has been the concept of *relativity*. State it in a simple way, the main argument we presented (chapters 2-4) is that a specific level of income is ample for one's standards of living whilst it is scant for another's; in economic theory terms: *preferences* vary between and within individuals, countries, cultures, and generations.

The starting point of economic theory lies on the assumption of fixed preferences; especially when economists probe fertility: " 'Becker's 'Chicago School' felt that there was no justification for assuming changes in preferences, and indeed felt that such an assumption was antithetical to the economic approach to fertility." (Macunovich, 1998: 54). Easterlin has been the pioneer on breaking up this doctrine: "Fixed preferences went first" (Easterlin, 2004: 14). The latter came not without any hesitations: "It is hard to overcome the preconceptions indoctrinated by graduate economics training. In the early years of my career, I sought faithfully to explain childbearing behavior on the basis of income and prices and to eschew appeal to preferences." (Easterlin, 2004: 13). By breaking up to economics doctrine bonds, Easterlin figured out the demographic behaviour in the post-World War II period: “By recognizing the role of changing material aspirations (preferences) along with growth of income, I was able to arrive at a plausible interpretation of the baby boom and bust – one consistent with the evidence” (Easterlin, 2004: 14).

The above might reveal why Easterlin is most recognized for his contribution to economics of happiness (“Easterlin paradox”) rather than his studies on fertility. One could hardly find any recent implementations of his relative income hypothesis out of the happiness literature. We think that, despite difficulties on employing the appropriate data, and a variety of approaches on aspirations, Easterlin’s relative income hypothesis – and overall his relativity approach – could provide us some explications on various demographic movements. This thesis has been such an attempt.

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