Fertility Transitions in Developing Countries: Convergence, Timing, and Causes*

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Abstract

This paper studies the dynamics of fertility in 180 countries in the period 1950–2015 and investigates the determinants of the onset of fertility transitions. We find evidence of convergence in three groups of countries, and distinguish the transitioning countries from those not transitioning. The estimation of the year of onset of the fertility transition is followed by an econometric analysis of the causes of this event. Instrumental-variable estimates show that increasing female education and reduced infant mortality are important determinants of fertility decline, while per-capita GDP has probably worked in the opposite direction. These results are confirmed by the application of Lewbel's (2012) methods where identification is based on heteroskedasticity.

Jel codes: J11, J13, I15, I25, C26

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1 Introduction

After World War II a demographic transition from high to low fertility occurred in many developing countries. The data from the World Population Prospects of the United Nations say that in the years 1950–1955, 140 of 201 countries (65.7% of the world population) had fertility rates greater than five children per woman. Sixty years later, fecundity is this high in only 22 of 201 countries (8.5% of the world population). Hence, since World War II, a large part of the world has adopted the advanced countries' reproduction model. An important feature of this phenomenon is the heterogeneity of the time paths of fertility across countries. Indeed, the timing of the onset of the transition varies substantially across countries, and in some cases fertility declines along non-monotonic time paths with trend inversions due to specific events that, however, do not change the direction of the long-term transition. Despite the important research produced in the last decades, the debate on the causes of the demographic transition is still open from the point of view of theory and empirical analysis (see Galor, 2011; Greenwood et al., 2017).

Our aim with this paper is to enhance understanding of the causes of demographic transitions. We investigate empirically the main stylised facts and determinants of the fertility transitions that occurred in many developing countries after World War II. Specifically, we make four main contributions. First, we study convergence and clustering of fertility transitions applying econometric methods that accommodate non-linearity and heterogeneity. Group convergence analysis allows us to pick out those countries that experienced a fertility transition. Second, we focus on the onset of fertility transitions and estimate the onset year. Third, we investigate the causal effects of income, infant mortality, and education, on the probability of entering a transition applying instrumental variables (IV) methods and verifying the relevance of the instruments using recent advances in testing the weak-IV hypothesis in models with multiple endogenous regressors. Fourth, we tackle the same causal inference problem using the Lewbel's (2012) alternative approach where identification relies not on exclusion restrictions, but on heteroskedasticity.

In the framework of this paper, the analysis of convergence provides the most important stylised facts. These derive from to the annual time series of fertility rates of 180 countries from 1950 to 2015. The application of Phillips and Sul's (2007) methods highlights the persistence of different models of reproduction. Although, in many countries today, fertility rates are low and close to those in the lowest countries, this process of convergence has not finished yet; in some cases, it has not even started. Hence, the end of the demographic transition is not imminent. Our analysis highlights three convergence groups. The first includes 14 countries – mostly in Sub-Saharan Africa – that provide no evidence of the beginning of a fertility transition. Today, in these countries TFR still remains close to six children per woman. The second group includes 25 countries showing a recent feeble declining TFR trends. The third group in our analysis includes 141 countries with the lowest TFR in the world, on average. In this group, two subgroups can be distinguished. One subgroup (59 countries) is characterised by the early onset of a fertility transition before the 1950s, while the 82 countries in the other group had high TFR in 1950 and later caught up with the low-fertility countries.

We concentrate on the group of transitioning countries to estimate the determinants of the probability of entering the transition. According to demographers (e.g., Chesnais, 1992; Kirk, 1996) and economists (e.g., Galor and Weil, 2000; Cervellati and Sunde, 2015), the crucial event in a demographic transition is the onset of a declining trend in the fertility time path. In economic theory, the onset of fertility decline occurs during the economy's transition from stagnation to sustained growth. While periods of declining birth trends were not rare in the past, they do not always identify a demographic transition characterised by the clear distinction between pre- and post-transition regimes. For these reasons, we investigate the causes of fertility transitions from the point of view of the onset timing.[1] We estimate the onset year of the fertility transition by applying the recent econometric methods proposed by Pierre Perron and his co-authors (Perron and Zhu, 2005; Perron and Yabu, 2009). These methods provide an estimation of the

[1] Here, we follow a recent strand of studies of the empirics of economic growth that concentrate on the determinants of growth events (e.g., Hausmann et al., 2005). The same approach has been followed by Spolaore and Wacziarg (2016) and Delventhal et al. (2018) to study the role of social interactions in the demographic transition.

date of a change in a linear trend's slope such that the trend function is joined at the time of break and the error in the model variable can be either stationary or have a unit root. This approach is quite new; existing estimates of the onset date of a demographic transition derive from ad hoc non inferential procedures based on the rate of change of fertility (e.g., Reher, 2004).[2]

We study the determinants of the onset of fertility transitions in 75 developing countries by estimating regression equations for two dependent variables: a binary indicator of transition for each country-period and the cross-country share of the number of years a country spent in a fertility transition in the number of years in the period under consideration. Our investigation focuses on the main factors that have been debated in the literature: income, education, and child mortality (see Schultz, 1997; Galor, 2011; Greenwood et al., 2017). We also control for important factors often discussed in studies of the fertility decline in the developing world: the rate of urbanisation, the Catholic and Muslim fractions of the population, the year in which women acquired the right to vote, and the legality of abortion. These variables approximate social and cultural factors relevant to parental reproductive behaviour.

One of the most important challenges to identifying causal effects in fertility models is the endogeneity of the variables commonly used to capture the main explanatory factors listed above. To tackle this issue, we apply IV methods. We instrument per-capita income with two variables. One is the trade-weighted world income introduced in the literature by Acemoglu et al. (2008) to capture the transmission of business cycles from one country to another through trade. The second is the measure of shocks to the country-level international oil price (Brueckner and Schwandt, 2014). To instrument infant mortality, we follow Acemoglu and Johnson's (2007) investigation of the causal effect of life expectancy on economic growth. They construct an instrument – predicted mortality – that approximates the strength of the international epidemiological transition occurring worldwide in the 1940s and 1950s. The instrument for the years of primary

[2] Quite independently, Delventhal et al. (2018) estimate the onset year of the demographic transition using a model that assumes three stages. In the pre-transition stage, fertility is constant. This stage is followed by a regime in which fertility declines linearly. In the final stage the birth rate is stationary again. The econometric framework is similar to the one used in this paper.

school of women of reproductive ages, 15–39 years, is the average educational attainment of adults aged 40–64 years. This variable is meant to approximate the human capital of teachers and parents in the production function of education.

Our main finding is that both increasing female human capital and declining infant mortality are important determinants of the onset of fertility decline in developing countries. In particular, on average, the probability of the onset of a fertility transition increases by about four to five percentage points if the years of primary schooling of women of reproductive ages grow by ten percentage points, while the same probability increases by about three to four percentage points when infant mortality decreases by ten percentage points. We find similar results when we apply Lewbel's (2012) method. Indeed, in this case, if women's primary school attainment increases by 10%, the onset year of a fertility transition comes 10.17 months earlier than without such increase, while a 10% reduction in infant mortality implies the same year comes around 7.19 months earlier. According to our estimates, income had either no role in triggering a transition or, more likely, worked in the opposite direction. Hence, our estimates are consistent with the results of several papers that find children are normal goods in parent preferences.

The next section discusses the related theoretical and empirical literature. Section 3 presents the convergence analysis of fertility time paths. In Section 4, we estimate the onset year of the fertility decline in transitioning countries. Section 5 presents the study of the determinants of the fertility transitions' timing. Section 6 concludes.

2 Related Literature

Several scholars of the fertility transition, demographers in particular, have stressed the importance of improved health status and declining mortality. The fact that, historically in several countries, mortality reductions preceded that of births provides a strong argument in support of this hypothesis (Kirk, 1996). The relationship between infant and child mortality and the choice of the number of children is the focus of a significant strand of the theoretical literature (e.g., Ben-Porath, 1976; Sah, 1991; Ehrlich and Lui,

1991; Boldrin and Jones, 2002; Kalemli-Ozcan, 2002; Soares, 2005; Doepke, 2005). In high-mortality environments, risk-averse parents may try to insure themselves against child death with more births. In this setting, a decline in infant mortality decreases uncertainty over child survival and reduces the demand for children. If the family's decision problem follows the sequence of births and deaths in the reproductive age, parents have an optimal desired number of surviving children and replace those dying in the early years of life with other births, until the targeted family size is reached. According to the replacement effect, decreasing infant mortality reduces total fertility without changing net fertility.

In one of the most important contributions to the economic analysis of fertility choice, the decline of births is explained by the rise in household income that occurred in developed countries (Becker, 1960; Becker and Lewis, 1973). In this model, a rise in income might induce parents to invest more on children quality, and this implies raising the opportunity cost of the number of children. The negative effect of income on fertility crucially depends on specific assumptions on parents' preferences.[3]

The influential Unified Growth Theory argues the demographic transition is caused by a technology-driven rise in the demand for human capital in the transition from stagnation to growth (Galor and Weil, 2000; Galor, 2011). The trade-off between quantity and quality of children in household choice explains the decline of fertility after the economic take-off. In this theoretical approach, another important factor of the onset of a demographic transition is the gender gap. According to Galor and Weil (1996), economic growth increases women's wages relative to men's wages because physical capital input is more complementary to mental labour than to physical labour, and women have a comparative advantage in mental labour. The decline in the gender wage gap causes the decline of fertility.

The empirical analysis of the demographic transition traditionally proceeds along two main lines. One describes the phenomenon, the other looks for the determinants. Descriptive studies concentrate on the time path of cross-country fertility, searching for

[3] Galor (2011) assesses this approach.

regularities, and find remarkable variation across nations of both the onset and the pace of fertility decline (e.g., Bongaarts and Watkins, 1996). Dorius (2008) uses the empirics of economic growth to investigate β-convergence and σ-convergence of the TFR of a panel of 195 countries from 1955 to 2005 and finds some evidence for fertility convergence in the world only after the 1980s. Strulik and Vollmer (2015) study the evolution of the world's fertility distribution from 1950 to 2005 by estimating the coefficients of a mixture of two normal distributions for each five-year interval. Silverman's test suggests that the number of peaks is not greater than two, and the TFR distribution is characterized by two peaks before 1995 and one peak after 1995.

In the literature on the determinants of fertility, the question of the identification of causal effects has been tackled quite recently and remains open. The endogeneity of child mortality has caught the attention of several scholars. Schultz (1997) instruments child mortality with per-capita calorie consumption and finds a positive and significant effect on TFR. The validity of this instrument can be questioned because it can be argued it may have direct effects on fertility. Acemoglu and Johnson (2007) make an important contribution to the identification of the effects of health improvements on economic growth and population dynamics with the careful construction of an instrumental variable, predicted mortality, that summarizes the exogenous components of the international epidemiological transition that occurred in the world between the 1940s and the 1950s. In this period, dramatic health improvements were determined by fundamental innovations in the chemical and pharmaceutical sectors (e.g., penicillin, new vaccines, DDT[4]) and by the adoption of a new international health policy with the establishment of the World Health Organization. Predicted mortality is used to instrument life expectancy in IV estimates of the long differences (1940–1980 or 1940–2000) of log per-capita income and log total births. Acemoglu and Johnson find a positive relation between life expectancy and births.[5]

Investigating the same research problem, Lorentzen et al. (2008) instrument both adult and infant mortality with the Malaria Ecology Index (Kiszewski et al., 2004). This

[4] Dichlorodiphenyl trichloroethylene.

[5] See also Hansen and Lønstrup (2015).

index has been derived with the aim of measuring the basic potential transmission of the disease at regional level in the world. Lorentzen et al. (2008) use this variable and other geographic and climate indicators to instrument adult mortality, infant mortality, secondary school enrolment, in estimates of the cross-country 1960–2000 average of the rate of fertility. Both mortality indicators display positive coefficients while the effect of school enrolment is negative. [6]

Although income and education enter most of the regression models we have discussed in this section, only a few studies concentrate on the search for valid and strong IV for them. Brueckner and Schwandt (2015) estimate the causal impact of income on population growth for a panel of 139 countries applying IV methods. Their identification strategy is based on the exogeneity of country-specific shocks to international oil prices on per-capita GDP. In their paper, the change in the international oil price multiplied by countries' average GDP shares of net oil exports instruments per-capita GDP growth in a panel regression of fertility change. Brueckner and Schwandt (2015) provide several arguments for the validity and relevance of their IV, and find a positive effect of income on fertility. Scotese Lehr's (2009) econometric analysis of net reproduction rates applies GMM methods to regression equations where education levels and enrolment rates are instrumented with the ratios of the numbers of telephones, radios, and newspapers to adults in the population. Scotese Lehr finds secondary enrolment rates have a negative effect on net fertility. Her paper does not address the relevance of the IV.

Murtin (2013) studies the determinants of the demographic transition on an international panel in the period 1870–2000. He estimates dynamic models of birth rate and infant mortality using Blundell and Bond's (1998) system-GMM estimator. Murtin advances the econometric research on the determinants of fertility assuming the endogeneity of the main factors, income, education, and mortality. As in many applications of the system-GMM technique, the author uses lags of the explanatory variables as internal instruments. The data's long time range allows the use of long lags in the IV definition.

[6] Recently, McCord et al. (2017) construct a time-varying version of the Malaria Ecology Index and use it to instrument child mortality in TFR panel regressions. They find a positive effect of child mortality on fertility.

Murtin (2013) finds that primary education is the only robust determinant of fertility, while education and income explain infant mortality. Although the econometric methodology employed in this study has found large success in the analysis of short dynamic panel data, the most recent research highlights some problems in system GMM that may hinder the identification of causal effects. Indeed, Bun and Windmeijer (2010) highlight a weak-instruments problem in the application of system GMM to short panels when data are time persistent and the variance of individual effects is larger than the variance of idiosyncratic shocks. This is the case of many cross-country panel datasets with short time dimensions. Bazzi and Clemens (2013) extend the analysis of the AR(1) model in Bun and Windmeijer (2010) to the case of multiple endogenous explanatory variables. Their simulations confirm the negative performance of the system GMM estimator in terms of bias. Murtin (2013) does not address the question of the relevance of the IV he uses in his dynamic panel estimates. This question remains still largely open in the comparative study of the determinants of the demographic transition.

3 Modelling Fertility-Rate Transition Paths and Convergence

In this section, we investigate the time paths of fertility in a large number of countries aiming to find the most important regularities while allowing for general features of the panel time series. Indeed, as argued above, the dynamics of fertility across the world in the years after 1950 are characterized by strong heterogeneity.

Recently, Phillips and Sul (2007) proposed a test of convergence assuming the variable X_{it} follows a factor model for individual i at time t:

$$X_{it} = \delta_{it}\mu_t, \tag{1}$$

where unobserved common factors μ_t and time-varying idiosyncratic components δ_{it} can be distinguished. In this model, the distance between the common factor and the systematic part of X_{it} is specific of the individual i and is not constant over time.

Convergence among the series X_{it} is defined as the long-run equilibrium of their ratios:

$$\lim_{k\to\infty} \frac{X_{it+k}}{X_{jt+k}} = 1 \text{ for all } i \text{ and } j. \tag{2}$$

Hence, relative convergence is equivalent to: $\lim_{k\to\infty} \delta_{it+k} = \delta$. This definition of convergence allows the analysis of time series which do not cointegrate although they follow the same stochastic trend in the long run. The *log t* test of convergence is defined in terms of the relative transition coefficients,

$$h_{it} = \frac{X_{it}}{\frac{1}{N}\sum_{i=1}^{N} X_{it}} = \frac{\delta_{it}}{\frac{1}{N}\sum_{i=1}^{N} \delta_{it}}, \tag{3}$$

which remove the common factor μ_t. Convergence now implies h_{it} is asymptotically equal to one, and the cross-sectional variance,

$$H_t = \frac{1}{N}\sum_{i=1}^{N} (h_{it} - 1)^2, \tag{4}$$

converges to zero. The test is based on the cross-sectional variance. The dynamics of the ratio $\frac{H_1}{H_t}$ over time are modelled by the regression equation

$$\log\left(\frac{H_1}{H_t}\right) - 2\log L(t) = a + \gamma \log(t) + \hat{u}_t, \tag{5}$$

where $L(t)$ is a slowly varying function: e.g. $L(t) = \log(t+1)$. Under the null, the cross-sectional variance tends to zero as t goes to infinity, meaning that $\log(H_1/H_t)$ diverges to ∞. In the case of divergence, H_t converges to a positive value and the left-hand side of the regression equation (5) diverges to $-\infty$. We perform a one-sided test of the null hypothesis of convergence ($\gamma \geqslant 0$)) using the t-statistic calculated using the estimate of γ and a heteroskedasticity- and autocorrelation-consistent (HAC) standard error. This statistic is asymptotically distributed as a standard normal. Appendix A provides more detail on the *log t* test.

Phillips and Sul (2007) propose a procedure to identify clusters of units where convergence occurs in the long run. The algorithm starts with the ordering of all units in the

panel according to the value of the last available observation, X_{iT}. Then, a core subgroup of units can be detected and the *log t* test can be used to assess whether another unit belongs to the core group or not. This procedure can be iterated until every unit in the panel is classified. In the following subsection, we apply this clustering method to the fertility rates in a panel of 180 countries.

3.1 Group Fertility-Rate Convergence

Data refer to the time series of 180 countries from 1950 to 2015.[7] The main source of the data is the World Bank's (2017) World Development Indicators (WDI). These annual time series are not available before 1960. To perform the *log t* test, we use as the first-year observation the estimate of TFR of the year 1950 drawn from United Nations World Population Prospects, 2017 Revision. Since Phillips and Sul (2007) recommend discarding the first 30% of the time series observations, we estimate regression (5) using panel data from 1970 to 2015. They also suggest filtering the series to remove possible business-cycle components in the TFR data. We employ Hodrick-Prescott filter for this purpose.

The initial approach to convergence analysis is the estimation of the regression (5) on data of the full set of 180 countries:

$$\log\left(\frac{H_1}{H_t}\right) - 2\log\left[\log\left(t+1\right)\right] = \underset{(2.25)}{0.609} \underset{(-14.52)}{-1.092} \log\left(t\right),$$

where t-statistics based on Newey-West standard errors are in parentheses. The null of convergence in fertility rates around the world is clearly rejected. The same result emerges from the regression we run on a more recent period, 1980–2015, to allow for more countries where the demographic transition has already begun. Indeed, the *log t* test provides $\hat{\gamma} = -0.92$ and $t = -9.98$. Given the remarkable generality and power of the *log t* test, this result strongly supports the view that the demographic transition is still an open question in many developing countries. The alternative to global convergence could be the case where every country has a distinct long-term TFR value, or, more

[7] Table A1 in on-line Appenidx A provides a lists of the countries in the convergence analysis.

realistically, the case of convergence within groups. We have pursued the investigation of such an alternative configuration of fertility dynamics by applying Phillips and Sul's (2007) clustering algorithm to the TFR data, and have found three groups (see online Appendix A.2 for details).

The first group (Group H) includes 14 countries converging to a high fertility rate. The *log t* test for the whole Group H provides a point estimate of $\hat{\gamma} = -0.091$, and a test statistic of $t_b = -0.57$. Table 1 shows the composition of each convergence group. Countries in Group H are characterized by persistently high fertility: On average their TFR was 6.53 in 1950 and 5.82 in 2015. Several countries of Sub-Saharan Africa are members of this group. With the application of the clustering algorithm we identify another convergence group of 25 countries. For this group, the *log t* test does not reject the null ($t_b = -0.40$). This group is composed of countries that, during the period, began a demographic transition that is far from concluding, given the average TFR greater than four births per woman (4.63) at the end of the period. We denote this as Group I to hint to the transition from high to intermediate fertility rate. The last convergence group derives from running the convergence test on the time series of the remaining countries. In this case, the *log t* test does not reject the null of convergence ($t_b = -0.23$), although its negative value indicates slow convergence. This Group L consists of 141 countries with low TFR at the end of the period; it includes all countries where the onset of the fertility transition occurred before 1950 and several countries where the same process started after that year.[8]

The characterization of the demographic transition in developing countries can be improved by taking advantage of the classification of countries in these three groups. Indeed, if we consider the fertility rate of countries in Group H in 1950 as typical of countries before the onset of the fertility transition, then we can reasonably classify as transitioning those countries that, starting from TFR equal to or greater than the minimum 1950 TFR in Group H, end in a different group during the period 1950–2015.

[8] Following Phillips and Sul (2009), we investigate the case for merging groups. The *log t* test does not suggest the merging of Group H and Group I since it gives the statistic $t_b = -20.94$.

Since the country in Group H with the lowest fertility in 1950, The Gambia, has TFR = 5.23, it seems clear that every country in Group I during the last six decades has been involved in a slow process of transition from TFR greater than 5.23 to lower, but still high, fertility rate (average TFR in 2015=4.63).

The most impressive part of the phenomenon under investigation is represented by the large number of nations that saw a strong process of fertility reduction and caught up with the low-fertility countries. Applying this classification criterion, countries in Group L can be clustered as 82 countries (Group H-L) that had TFR in the range of Group H at the beginning of the period, and 59 countries starting with low fertility (Group L-L). Some Sub-Saharan countries (Botswana, South Africa, Swaziland, Zimbabwe) enter Group H-L which also includes many other countries of Latin America, Asia, Africa, and Oceania. Table 1 summarizes the TFR dynamics in the five clusters of this analysis.

Figure 1 depicts the average TFR over the period for each group, showing some interesting patterns. Indeed, in the early years of the period all groups, excluding Group L-L, display very similar fertility rates in the range 5.23–8. Afterwards, each group shows a distinctive TFR pattern. The line of fertility in Group H is located at the top of Figure 1. In these countries, the number of children per woman follows an increasing trend that quite recently becomes decreasing, ending in 2015 with a still-high TFR, 5.82. The last trend seems too recent and weak to be considered the signal of the onset of a demographic transition. The line at the bottom of Figure 1 depicts the fertility rate of those countries that underwent a demographic transition before the 1950s. Their TFR in 1950 is already low, around three. Later, reproductive behavior in these countries steadily changes, and the fertility rate declines below the significant value of two. Hence, for opposite reasons, fertility dynamics in countries in Group H and in Group L-L do not identify the onset of a transition from high to low TFR after World War II.

But this is not the case for the other countries we consider in our convergence analysis. Indeed, the fertility paths of many countries of Group I and Group H-L are characterized by an initially high and increasing TFR followed by a decreasing path that leads to different TFR values. The countries of Group H-L exhibit a stronger change in reproductive

behavior. Indeed, Figure 1 shows that in this group, on average, the onset of the fertility decline occurs during the sixties, before other transitioning countries, and then it proceeds along a steep path. At the end of the period, average Group H-L TFR approaches that of Group L-L. The curve depicting the dynamics of TFR of countries in Groups I displays the same shape as that of Group H-L, but the declining part of the curve has a more gradual slope and at the end of the period fertility is still significantly higher than that of Group L.

The descriptive study of fertility transitions raises the question of the forces that have driven the decline in fertility. The answer to this question would have important implications for the empirical assessment of the theory of economic development and for population policy in developing countries. We pursue this goal in the next sections.

4 The Onset Year of Fertility Transitions

Most of the TFR time series in those countries that experienced a demographic transition clearly show a pattern characterized by two distinct trends: before a break date the trend is constant or slightly increasing; after that date the trend changes and becomes decreasing. To search for the determinants of this phenomenon, we concentrate our analysis on the countries in Group I and Group H-L. Our empirical strategy consists of the estimation of the year of the onset of the fertility decline, and the econometric analysis of the factors that may explain the onset of the demographic transition.

In this section, we focus on the annual TFR time series of 107 countries we have classified into Groups I and H-L. From 1960 to 2015, the data are those we have already used in the convergence analysis. Annual time series of fertility from 1950 to 1959 are the estimate of TFR drawn from United Nations World Population Prospects, 2017 Revision.

To estimate the onset year of the fertility transition, we rely on the following model of the trend of a generic variable y_t.

$$y_t = \mu + \beta_1 t + \beta_b B_t + u_t, \qquad t = 1, ..., T \tag{6}$$

where u_t is a random component and B_t is a dummy variable for a one-time change in the slope of the trend at the year T_1:

$$B_t = \begin{cases} 0 \quad if \quad t \leq T_1, \\ t - T_1 \quad if \quad t > T_1. \end{cases}$$

Note that the trend function in this model remains joined at the year of the break. Otherwise, the trend would verify a discontinuity at T_1. Actually, the demographic transition is usually thought of as the outcome of a process that gradually changes the behavior of households, and can hardly be meant as the result of some sudden shock to the economy. In the econometric application of this model, we follow the recent research by Pierre Perron and his coauthors. In particular, Perron and Zhu (2005) analyze the consistency of the coefficient estimates of model (6) under quite general assumptions that allow either stationarity or a unit root in u_t. For each break date T_1 the application of OLS to the equation (6) provides estimates of the coefficients μ, β_1, β_b. Among all the admissible dates, the break date that minimizes the corresponding OLS value of the sum of squared residuals is the estimate of T_1. Perron and Zhu (2005) prove that this estimate is consistent and its distribution converges to a normal that does not depend on the nature of the serial correlation of the model's error component. This methodology assumes that the trend function (6) includes one break in the slope. Given a break date, a test of structural change of the model (6) at that date can be performed by applying Perron and Yabu's (2009) methods. The null hypothesis of the test is $\beta_b = \beta_2 - \beta_1 = 0$, meaning the trend function (6) does not change in the period under consideration. The Wald test has the limit distribution, $\chi^2(1)$. Details on this test can be found in Appendix B.

We apply this econometric framework to the logarithm of the annual TFR series of those countries that underwent a demographic transition after 1950. As in Perron and Zhu (2005), the estimation of the coefficients of (6) is limited to the range of the time series defined as $\varepsilon \leq \lambda_1 \leq 1 - \varepsilon$, where $\lambda_1 = T_1/T$, and ε is a trimming parameter. In this paper, we set ε to 10% and search for a break date between 1956 and 2009.

Table B1 in online Appendix B presents the results of the econometric analysis. For each country the Table shows the estimate of the year of the onset of the fertility transition, the slope coefficients, β_1 and $\beta_2 = \beta_1 + \beta_b$, and the Wald test for a break in the slope of the trend, W_{RQF}. For each country, the plot of the time series of TFR is shown in Appendix B. In each graph, a vertical line represents the estimate of the onset year.

The Wald test confirms that the trend in fertility changed significantly in each country we selected. Many countries show estimates of the test statistics with values much greater than the usual threshold at the 1% significance level. Most of the countries in Group I display an estimated break year in the 1970s and even later. The average rate of decline of fertility after the onset is close to 2% per year. The last two results seem to suggest that in these countries' TFRs can be further reduced. The strongly transitioning countries of Group H-L are often characterized by an onset year in the 1960s and a significant rate of reduction of fertility in the range 2%–5% per year. The estimated onset year of fertility decline in China is 1963, but the test statistics W_{RQF} is close to zero, although in this country TFR declines from 6.6 in 1950 to 1.6 in 2015. The plot of China's fertility rate since 1950 highlights three periods with increasing fertility and three periods of declining fertility. The changing phases of China's fertility dynamics could be responsible for the low value of W_{RQF} statistic for $T_1 = 1963$. However, the graph suggests 1963 can be reasonably considered the year that marks the onset of the transition from high to low fertility in China.[9]

5 Determinants of Fertility Transition Onsets

Now we have a quantitative description of the fertility decline occurring in many countries after World War II. According to economists and demographers, the explanation of the onset of this phenomenon is crucial for understanding the main forces driving the demographic transition. The literature we reviewed in Section 2 converges on the following main causes: income, mortality and health, education, and technological pro-

[9] Reher (2004) estimates the onset of fertility decline in China in 1970, using quinquennial data on births.

gress. Furthermore, the most recent research argues a significant role for the evolution of gender differences. Here, our aim is the investigation of the main factors that triggered the fertility transition in those countries where it occurred between 1956 and 2009. Two indicators of the transition will be used as dependent variables in regression analysis. The first is a binary indicator, $onset_{it}$, that takes value zero if the period t precedes T_{1i}, the year of the onset of a fertility decline in country i, and a value one otherwise. The second dependent variable, $transition_i$, measures the portion of the period under consideration during which country i experienced a fertility transition, and is simply the average of $onset_{it}$ taken over the time dimension. This variable is continuous in the interval $(0, 1)$ and this property allows the application of Lewbel's (2012) heteroskedasticity-based identification methods. We follow the main strand of the literature and focus on the likely causes discussed in Section 2.

5.1 Econometric Model and Data

The econometric model we use to investigate the determinants of the probability of the onset of the fertility transition is specified as:

$$onset_{it} = \alpha \log(GDP_{it}) + \beta \log(infant\ mortality_{it}) + \gamma \log(schoolw\_15-39_{it}) + \quad (7)$$
$$+ \mathbf{x}_{it}\boldsymbol{\delta} + \mu_t + c_i + u_{it}, \quad i = 1, ..., N, \quad t = 1, ...T,$$

where GDP is the real per-capita GDP, $infant\ mortality$ is the infant mortality rate, $schoolw\_15-39$ denotes the average years of primary school of women in the 15–39 age range. Several studies find that primary education is particularly effective in explaining fertility (e.g., Murtin, 2013). The variable $schoolw\_15-39$ stresses the role of women's human capital in the choice of fertility (Galor and Weil, 1996). $\mathbf{x}$ is a vector of time-varying and time-invariant control variables, μ_t is the time effect, c_i is the country effect, which can be assumed correlated or uncorrelated with the random error u_{it}. The time

index t refers to nine five-year nonoverlapping periods from 1955–1959 to 1995–1999.[10] In particular, the variable *onset* has been constructed as described above but with reference to five-year periods. Indeed, the benchmark period is defined by the interval that includes the year of the onset estimated in Section 4. Hence, *onset* takes value zero if t precedes the time interval during which a transition started, and value one otherwise. In this way, measurement error in the dependent variable *onset* should be smaller. The time effects imply that our model accounts for nonlinear trends common to all countries in the sample.

5.1.1 Data

The dataset comes from 75 countries with complete time series on per-capita income, educational attainment, and infant mortality. The source of real per-capita GDP is the Maddison Project at the Groningen Growth and Development Centre. Data on the average years of primary school of women in the 15–39 age range come from Barro and Lee (2013). The infant mortality time series is from World Population Prospects: 2017 Revision of the United Nations Department of Economic and Social Affairs, Population Division.

Five control variables that approximate for social and institutional factors complete the model specification. Each country's urbanization is measured by the percentage of population at mid-year residing in urban areas (*urbanization*), and the data are from the U.N. World Urbanization Prospects: 2014 Revision. The data on the Catholic (*Catholic*) and Muslim (*Muslim*) fractions of population in 1970 and 2000 are from the Religion Adherence dataset used by Barro and McCleary (2003).

The abortion index (*abortion*) of Bloom et al. (2009) measures the legal availability of abortion. The index takes integer values in the range 0–7. For each of seven reasons each country receives a score of 1 if abortion is allowed, while if it is not permitted the score is 0. The index is constructed as the summation of the scores for the seven reasons.

[10] Note that the estimate of T_1 is an year before 2000 for all the transitioning countries but Afghanistan. Afghanistan is not considered in the analysis of this section due to missing data. Hence, the dependent variable $onset_{it}$ after 2000 would take the value 1 for all countries in the sample.

Annual data are available for the period 1960-2006 and in estimates we use the data of the first year in each of the five-year nonoverlapping periods from 1960-1964 to 1995-1999 (e.g., 1960 for the period 1960-1964 and so on).

A recent strand of the literature on the demographic transition (e.g., Diebolt and Perrin, 2013; Greenwood et al., 2017) argues for a significant role of the empowerment of women in the onset of the transition. In our regression analysis, we account for this phenomenon with a variable (*women vote*) that records the year in which women acquired the rights to vote and to stand for election. The source of the data is the Inter-Parliamentary Union. Online Appendix C1 provides a description of the variables used in this section. Summary statistics are in Table 2, and the correlation matrix of the time-varying variables is shown in Table 3.

Income, infant health, and mother's human capital can hardly be assumed exogenous to fertility dynamics. Endogeneity may derive from general equilibrium models of growth and the demographic transition in the form of reverse causation. Endogeneity could also be the consequence of omitted variables and measurement errors. As already argued, the joint econometric evaluation of the most important determinants of the demographic transition is one of this paper's main challenges. We pursue this objective using the IV methodology.

The traditional approach of simultaneous equations has been deeply criticized, and present econometric theory stresses the difficulties in deriving reliable causality statements from the analysis of models with endogenous regressors (e.g., Angrist and Pischke, 2009). In this respect, the IV method offers a well-developed framework, although the requirements the IV must meet can be quite rigorous. Indeed, IV must be correlated with the endogenous regressors but uncorrelated with the random error in the structural equation. Hence, a valid instrument has no direct relation with the dependent variable in the structural equation. The approach of several papers that analyze cross-country econometric models, and in particular of those dealing with the determinants of population growth (see, e.g., Acemoglu and Johnson, 2007; Cervellati and Sunde, 2011; Brueckner and Schwandt, 2015; Hansen and Lønstrup, 2015), relies on the estimation of models with

one endogenous explanatory variable. These papers use IV whose validity and relevance have been carefully argued.

However, it must be noted that the orthogonality condition between IV and the error term can be violated when the estimated equation include no relevant variables correlated with the instruments. This problem affects several papers that investigate macroeconomic models with cross-country data, as shown in Bazzi and Clemens (2013). One solution for this difficulty is the specification of models with multiple endogenous regressors. On one hand, this approach is useful to purge the correlation between IV and endogenous regressors from the correlation between instruments and random error. On the other hand, it brings about the significant difficulties that characterize the estimation of models with multiple endogenous regressors. Indeed, in this case, identification needs more valid instruments, and relations among them are very important. In the following, we present the estimates of both equations that include one endogenous explanatory variable and equations that include three endogenous regressors.

5.1.2 Choice of Instrumental Variables

Recent research on demographic change and economic growth proposes careful IV construction that can be used in the context of this analysis too. Acemoglu and Johnson (2007) estimate the effect of life expectancy on growth and births using an IV (*predicted mortality*) that captures the dramatic reduction in the mortality from infectious diseases occurred in the world in the 1940s and 1950s as a consequence of three crucial events: the diffusion of important innovations in chemicals and drugs (penicillin, streptomycin, DDT) the development of new vaccines, and the creation of the World Health Organization and the start of a new international policy for public health. The *predicted mortality* instrument of country i at time t is constructed as an index of the mortality rate from 15 infectious diseases at 1940 and the following decades. For each disease, the index distinguishes mortality before the global intervention and after. In the context of our sample, data on predicted mortality are available for 42 countries. Here, the variable *predicted mortality* refers to the original data lagged 20 years.

Brueckner and Schwandt, (2015) estimate the casual effect of per-capita income's rate of growth on population growth. Their main instrument is a country-specific oil price shock variable: "the change in the log of the international oil price weighted with countries' sample average net-export shares of oil in GDP" (Brueckner and Schwandt, 2015, p. 1654). This variable (*oil shock*) has permanent effects on the level of per-capita GDP, which justify its use as an instrument for per-capita income in this paper. The same paper uses as additional IV the trade-weighted world income proposed by Acemoglu et al. (2008). This variable (*world income*) captures the transmission of business cycles from one country to another through trade. Indeed, it measures how the income of country i is connected to its trading partners' income.[11]

In the literature on the causal effect of education on country-level outcomes, this variable is often instrumented by its lagged values (e.g., Murtin, 2013; Barro and Lee (2015); Barro, 2015). Here, fertility depends on the mother's human capital, approximated by the average years of primary schooling of women of reproductive age. In the productive process of schooling, one of the most important inputs is the human capital of teachers. This variable can be chosen as an instrument of women's education in equation (7) because it can hardly be argued that it affects fertility directly or through other variables not considered in (7). At an aggregate country level, average teachers' education can be approximated by the average educational attainment of the adult population. We construct this variable as the total average years of schooling of men and women aged 40–64 (*school_40–64*), drawing the data from Barro and Lee (2013).

5.2 Results

Our investigation of the determinants of the onset of fertility transitions starts with the estimation of econometric models that highlight some basic features of the relationships among the variables and will be a useful reference for the subsequent analysis. All the regressors are assumed exogenous. The estimates of standard errors are robust to hetero-

[11] Further details on the construction of this variable can be found in Acemoglu et. al., (2008).

skedasticity and clustered at the country level. Estimates of model (7), shown in Table 4, take country heterogeneity into account with the specification of fixed effects (Columns 1 and 2) and random effects (Columns 3–6). The most important explanatory variables support our choices with significant coefficients of the expected sign. The coefficient for per-capita income shows a negative sign, meaning that increasing income had a negative effect on the probability of the onset of a fertility transition. This result is consistent with the Beckerian view that children are a normal good in parents' preferences and confirms recent results from Lindo (2010), Black et al. (2013), and Brueckner and Schwandt (2015). Countries where women's educational attainment increased had a greater probability of undergoing a fertility transition, as did countries where infant mortality fell. Hence, the onset of fertility decline seems strongly associated with both female education growth and health improvement. It can be noted that the urbanization rate stands as one of the most important correlates of the onset of a fertility decline. However, OLS regressions do not enable us to establish causal relationships. We tackle this issue in the rest of this section.

5.2.1 IV Estimation: One Endogenous Regressor

In this section we present the results of the estimation of panel, fixed-effects, IV models of the dependent variable *onset* where the set of regressors includes in turn *GDP*, *infant mortality*, *schoolw_15–39*.[12] Regressions also include the time-varying variables *urbanization* and *abortion*. All the right-hand-side variables enter the estimates as log transformations. Table 5 presents the results of nine regressions with per-capita GDP endogenous. The first six columns refer to exactly identified equations where the instrument is *oil shock* in the first three columns, and *world income* in the next three. The last three columns present the coefficients of the same equation estimated using two IV: *oil shock* and *world income*. The results show estimates of *GDP* not significantly different from zero in five over six of the exactly identified models, while the estimates are negative and significant in the last three overidentified equations. The Anderson and Rubin's F test

[12] We used the unofficial STATA command ivreg2 by Baum, et al. (2010). See also Baum, et al. (2007).

of the null: The endogenous regressors in equation (7) are irrelevant, in eight over nine model estimates accepts the null. This test is robust to weak instruments.[13] The variable *urbanization* displays positive and significant coefficient estimates in all the specifications. In the first stage regression results of Table 5, the IV coefficients show significant estimates of the expected sign, and the Kleibergen–Paap Wald F-statistic (Kleibergen and Paap, 2006)[14] takes a value greater than 10 in all but one regressions. Hence, following the rule of thumb suggested by Staiger and Stock (1997), we do not accept the null of weak instruments. In the last three regressions, the Hansen J statistic takes values that support the validity of the orthogonality conditions.

The effect of infant mortality on the dependent variable *onset* is investigated in two sets of estimates. Columns 1–3 in Table 6 show the results of fixed-effects 2SLS estimates where the instrument is *predicted mortality*, whereas Columns 4–6 refer to estimates with two IV, *predicted mortality* and *predicted mortality*, interacted with a dummy variable for the Sub-Saharan African countries. The point estimates of infant mortality are all negative and statistically significant. The first-stage estimates in these 2SLS regressions highlight the relevance of the chosen instruments, with coefficients for *predicted mortality* statistically significant at the 1% level, and large values of the Kleibergen–Paap Wald F-statistic. The Hansen J statistic lends support to the assumption that the IV are not correlated with the random error. In all regressions, we reject the null of the Anderson–Rubin test at the 1% level. The estimates of the effect of infant mortality on the probability of a fertility transition range from -0.54 to -0.64.

The last application of the one-endogenous-regressor approach refers to women's educational attainment. Table 7 presents the results of six regressions: In the first three, we instrument log *schoolw_15–39* with log *school_40–64*, while in the last three, we use two IV: log *school_40–64* and log *school_40–64*, interacted with a dummy variable for the South Asian countries. In every specification the coefficients of log *schoolw_15–39* and

[13] Baum et al. (2007) and Andrews et. al. (2018) discuss this test.

[14] The use of the robust F statistic to test weak IV is not supported by econometric theory. Olea and Pflueger (2013) propose the Effective F statistic for models with one endogenous regressor. This test is robust to heteroskedasticity, serial correlation, and clustering. In the case of just identified models, both tests are equivalent.

log *urbanization* are strongly significant and positive, as expected. The same can be said of the coefficients of the IV in the first-stage estimates. Consistent with these results, the Kleibergen–Paap Wald F-statistic estimates are greater than 40. Once again, the Hansen J statistic lends support to the overidentifying-exclusion restrictions. The estimates of the effect of the educational attainment of women aged 15–39 on the probability of a fertility transition range from 0.33 to 0.47.

5.2.2 IV Estimation: Multiple Endogenous Regressors

The main problem with the estimation of separate regressions when theory suggests the relevance of multiple endogenous explanatory variables is the assumption that in each model the IV are not correlated with the omitted variables, which necessarily enter the error component. In this section, we estimate the regression equations with per-capita income, infant mortality, and women's educational attainment jointly in the right-hand side. As already argued in this paper, this strategy could produce improved estimates of the causal effects, but their validity depends on more restrictive identification requirements.

We estimate the model (7) applying 2SLS and limited-information maximum likelihood (LIML), which are efficient under the assumption of a homoskedastic error term, and two-step GMM, a method that is efficient to arbitrary heteroskedasticity. The recent literature on weak instruments argues for LIML's better performance in finite samples compared to 2SLS and two-step efficient GMM (Baum et al., 2007; Stock and Yogo, 2005). Table 8 presents the results of the fixed-effects panel regressions where the three endogenous regressors, log *GDP*, log *schoolw_15–39*, and log *infant mortality*, are instrumented by *world income*, log *school_40–64*, *predicted mortality*, and other variables obtained from the interactions between the same variables and time period or world region dummy variables.

Hence, the set of instruments includes the variables: *world income*, plus two variables obtained from the interactions between *world income* and the dummies for the period 1970–1974 and for the countries of South Asia, East Asia and Pacific; log *school_40–64*,

and its product with the dummies for the periods 1960-1964 and 1970-1974; *predicted mortality*, plus six variables obtained from the interactions between *predicted mortality* and the dummies for the periods 1960–1964 and 1965–1969, for Latin American countries, for Sub-Saharan African countries, for Europe and Central Asia, and for the Middle East and North Africa. The last six interactive instruments capture the distinctive impact of the international epidemiological transition on infant mortality in the 1960s and 1970s with respect to more recent decades, and account for regional differences in the same effect.

In Table 8, the 2SLS estimates confirm the results of the preceding regressions that highlight how growth of per-capita income acted to restrain fertility, while the education of reproductive-age women stands as one of the most important factors of the transition. The estimates of the effect of infant mortality also confirm those obtained in the single-endogenous-regressor approach. The coefficients show negative signs and become highly statistically significant when the estimation method admits the heteroskedasticity of the error term (GMM). The infant-mortality coefficient estimates assume values ranging from -0.30 to -0.47, lower than those obtained in the previous IV regressions, while the coefficients of log *schoolw_15–39* range from 0.38 to 0.53, with values very close to those obtained in the previous IV estimates.

Table C2 in online Appendix C presents the estimates of the coefficients of the IV in the first-stage equations. In each equation, the IV specific of the endogenous regressor display significant coefficients with the expected signs. To assess the validity of the IV used in the second-stage estimates, we test for underidentification of the regression model, weak instruments, and overidentifying restrictions. In Table 8, all the regressions display values of the Kleibergen–Paap rk Lagrange-multiplier test statistic that suggest the rejection of the null of underidentification at the 10% significance level. We test for weak instruments using the conditional first-stage F-statistic Sanderson and Windmeijer (2016) recently proposed for equations with multiple endogenous variables. Assuming the error term is i.i.d, this statistic can be used to test the null of weak IV in relation to a given maximum relative bias. Sanderson and Windmeijer (2016) argue that the

critical values tabulated in Stock and Yogo (2005) can be used to conduct the test with the conditional F-statistic. However, the assumption of i.i.d disturbances often seems unjustified, particularly in linear-probability models.[15] Accordingly, all the test statistics in our regressions are robust to heteroskedasticity. In this case, testing weak IV is usually conducted by applying Staiger and Stock's (1997) rule of thumb.

In our regressions, this rule suggest the reduced-form log *schoolw_15–39* is unaffected by the use of weak IV. Indeed, in this case the conditional F-statistic assumes values greater than 66. In the estimates of the reduced form of log *infant mortality*, the same statistic takes values greater than 10 in two specifications and close to 10 in the third (Sanderson-Windmeijer F-statistic=8.56). Identification of the reduced form log *GDP* seems more problematic. Indeed, the Sanderson-Windmeijer F-statistic takes value 10.69 in one model specification, and 6.17, 6.77 in other two. However, it can be noted that in the second-stage equations the three estimates of the coefficient of log *GDP* are almost identical. Further support to our estimates comes from Anderson and Rubin's F test because in all regressions, we reject the null of the Anderson–Rubin test at the 1% level. The Hansen J statistic suggests the validity of the orthogonality conditions in all regressions.

Recent research on the IV estimation method (Young, 2018; Andrews, 2018) highlights the negative influence of outliers and non-homoskedastic errors on the quality of inference, in particular when weak IV are used. In online Appendix C2, we show that outliers have little influence on our GMM estimates of the basic model (specification 7 in Table 8), and the null of normally distributed residuals cannot be rejected. We estimate the model removing one country in each regression. Table C3, panel *A* shows how the estimates of the coefficients of the three endogenous regressors remain largely significant in all but one sample that gives a p-value of 0.105 for the coefficient of log *infant mortality*. Similar, but less strong, evidence comes from panel *B* in Table C3 that summarizes the distribution of the estimate of the first-stage F-statistic. The equation of log *schoolw_15–39* remains always strongly identified, while 66% of the Fs from the estimates of log *GDP* and log

[15] Andrews et. al. (2018) discuss weak IV focusing on the case errors are non-homoskedastic.

infant mortality take values greater than 10, and 90% greater than 9.

The evidence on the determinants of a fertility-transition onset we gathered from the last set of estimates clearly confirms the results of the estimates of one-endogenous-regressor models on the role of female education and infant mortality and per-capita income. We carry out one further robustness check investigating fertility transitions by taking advantage of recent advances in the econometric identification of simultaneous equation systems that do not rely on exclusion restrictions.

5.2.3 Identification through Heteroskedasticity

The paucity of valid IV in the econometric modeling of the aggregate determinants of fertility is clearly evident in the literature. The alternative approach to identification proposed by Lewbel (2012) is based on the heteroskedasticity of the error terms and does not rely on exclusion restrictions. Indeed, Lewbel's (2012) method is based on IV generated from variables exogenous to the model. The application of this identification method to binary outcomes like *onset* is not clearly justified because the paper does not address this case, and the extension is not straightforward. Consequently, we construct a continuous outcome – *transition* – that, as argued above, has the same meaning of *onset* but in a single cross-sectional dimension of the data. In the following econometric model, the regressors are averages across time of the variables in model (7):

$$transition_i = \mathbf{y}_i\boldsymbol{\gamma}' + \mathbf{x}_i\boldsymbol{\delta}' + c + \epsilon_i, \qquad i = 1, ..., N, \tag{8}$$

where $\mathbf{y}$ is a vector of endogenous regressors. Identification rests on $\mathbf{z}$, a vector of exogenous variables that may be equal to $\mathbf{x}$ or may include other exogenous variables.[16] The heteroskedasticity of the error terms of the reduced-form equation of the endogenous regressors is a fundamental assumption here. If the variables in $\mathbf{z}$ are correlated with the square of the first-stage error terms but uncorrelated with the product of these errors and the structural equation error component, ϵ, then equation (8) is identified. The first of the

[16] Lewbel (2012) assumes that variables not in $\mathbf{x}$ may enter $\mathbf{z}$.The variables of $\mathbf{z}$ can be discrete or continuous.

last two assumptions can be tested using the Breusch–Pagan test of heteroskedasticity. Lewbel (2012) shows that, in the case of simultaneous equation systems, both assumptions are satisfied when the presence of an unobserved common factor brings about the correlation of errors across equations.[17] The estimation algorithm proceeds in two stages. In the first stage, we estimate the reduced forms of each endogenous regressor by OLS and save the residuals. The residuals are then used to generate the instruments for the second stage. Each instrument is the product of a variable in $\mathbf{z}$ centered at the mean and the residual. In the second stage, the structural equation is estimated by GMM using the generated instruments and the variables in $\mathbf{x}$ as IV.

The exogenous variables we use to generate Lewbel's IV are the log of the urbanization rate, a dummy variable for the countries of South Asia, a dummy for countries in the Middle East and North Africa, and the interaction term between the variable *Muslim* and a dummy for East Asia and Pacific countries. We add to the generated IV the share in current GDP of merchandise export at current PPP (*export*) as an external instrument for per-capita income. Openness to international trade has been used as an instrument for GDP in several papers. However, the correlation between the two variables in the sample of transitioning countries under our investigation is weak, while that between per-capita GDP and the share of exports is quite important. Data on this instrument are drawn from the Penn World Table version 9.0. Table C4 in online Appendix C presents the correlation matrix for the regressors.

Table 9 presents the results of the Lewbel's GMM estimates for *transition*. The endogenous regressors are the per-capita GDP, the average years of primary school of women in the 15–39 age range, and the log of the infant mortality rate. We estimate four different equations, starting from the basic – which includes only the endogenous regressors and the variables in $\mathbf{z}$ – and adding each time the variables *Catholic, Women vote*, and log *abortion*. The fundamental assumption of heteroskedastic error terms in the first-stage equations finds support in the Breusch–Pagan test: The null is rejected using

[17] The method of Lewbel (2012) has been applied in several papers including, more recently, Emran and Hou (2013); Millimet and Roy (2016).

the usual significance levels.

To compare the marginal effect of the logged variables with that of the untransformed variables, Table 9 presents in square brackets the semielasticity, the change of *transition* after a proportionate increase in the regressor, averaged across the countries in the sample. In the first two specifications the coefficient of *GDP* is statistically significant but in the other two the null that the coefficient is zero cannot be rejected. In every specification the coefficients of *schoolw_15–39* and log *infant mortality* are strongly significant. In the most complete specification, the coefficient of *schoolw_15–39* implies that if the educational attainment of women increases by 10% the year of the onset comes 10.17 months earlier, while the corresponding coefficients for infant mortality implies the event occurs 7.19 months earlier.[18]

The diagnostic tests support the use of the instruments based on heteroskedasticity. The Hansen J statistic implies the overidentifying restrictions cannot be rejected. The null that the estimated equations are underidentified can be rejected at the 1% confidence level using the Kleibergen–Paap rk statistic. The explanatory power of the Lewbel's (2012) instruments seems reasonably strong. Indeed, in the reduced forms, the conditional F-statistic takes values from 18 to 78. The Anderson–Rubin F test strongly rejects the irrelevance of the endogenous regressors in the structural equation. A basic assumption of the model is linearity. Table 9 presents the results of the application of the Ramsey RESET test proposed by Pesaran and Taylor (1999) for IV estimators.[19] In all regressions, the null hypothesis that there are no neglected nonlinearities can be comfortably accepted.

[18] We calculate the effect of a change in an explanatory variable x on the onset year T_1 considering that:

$transition = \frac{2009-T_1}{2009-1955}$. Indeed, in this case we have: $dT_1 = -54\beta\frac{dx}{x}$, where β is the coefficient of log x in a regression equation of $transition$.

[19] We used the unofficial STATA command *ivreset* by Mark E. Schaffer: https://ideas.repec.org/c/boc/bocode/s455701.html

6 Conclusions

This paper analyzes the fertility transition that has occurred in many developing countries since World War II from a new empirical perspective. Indeed, first, we investigate the main features of the time path of fertility with a general approach that accommodates heterogeneity and nonlinearity. Second, we study the causes of the event that distinguish the periods before and after the transition's onset.

The descriptive part of the analysis confirms that the important process of fertility decline occurred in many developing countries and highlights a process of international convergence of countries in three groups. Transitions from high- to low-fertility regimes are clearly characterized. Based on the descriptive analysis, the paper provides significant econometric evidence supporting the theoretical analyses that attribute the demographic transition to growth in women's educational attainment and to significant reductions in child mortality. Regression analysis also produces evidence consistent with the view that children are normal goods and increasing income brings about increasing fertility.

The results of the causal analysis derive from the application of IV methods to models with multiple endogenous regressors. Our use of new inferential tools to assess the quality of the instruments matches the choice of IV proposed in other influential papers. This issue is crucial for the analysis of macroeconomic panel data and represents an important area of investigation in both econometric theory and applied demographic economics. This paper suggests that reproductive behavior significantly depends on education and health. Our results advance the comparative study of the determinants of the demographic transition.

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