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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. The application of Phillips and Sul's (2007) test to fertility rates provides evidence of convergence in three groups of countries. This information is used to 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. The instruments for per-capita income are the trade-weighted world income (Acemoglu et al., 2008) and the oil price shock (Brueckner and Schwandt, 2015). Mortality is instrumented with the predicted mortality (Acemoglu and Johnson, 2007), and mothers' education with the years of schooling of people aged 40–64. These results are confirmed by the application of Lewbel's (2012) methods where identification is based on heteroskedasticity.

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1 Introduction

In the history of the modern world, fertility declines during the transition from an age of economic stagnation to an age of sustained growth driven by technological progress. In this evolutionary phenomenon, demographic change is also characterized by higher investments in human capital and increasing life expectancy. These stylized facts arise from the history of the advanced countries (Galor, 2011) where the onset of the demographic transition occurred between the late nineteenth century and the first decades of the twentieth century. In many developing countries, the decline of fertility began after World War II.

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 first significant distinction is between countries converging to the low-fertility group and countries that persist in the pretransitional regime. Even among the transitioning countries, the timing of the onset of the transition varies substantially. Moreover, in some cases, fertility declines along nonmonotonic 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 Guinnane, 2011; Galor, 2012). In particular, three main issues are under discussion. One issue concerns the effect of income on fertility, where recent evidence (e.g., Brueckner and Schwandt, 2015) suggests a positive sign. Thus, it seems unlikely that this variable triggered the transition. The second question concerns the identification of child mortality improvements as a determinant of the fertility decline, where both theory and evidence raise several doubts (e.g., Doepke, 2005; Murtin, 2013). The third issue relates to the explanatory power of the hypothesis that the demographic transition is triggered by the rise in the demand for education that occurs in the process of development, with both theoretical and empirical contributions (e.g., Galor and Weil, 2000; Scotese Lehr, 2009; Galor, 2011; Murtin, 2013).

We contribute to this debate with an empirical investigation of the main features and determinants of the fertility transitions that occurred in many developing countries after World War II. Specifically, we make four main contributions. First, we propose an empirical framework for the study of convergence and clustering of fertility transitions that accommodates nonlinearity and heterogeneity. Second, we estimate the fertility transitions' onset year. Third, we investigate the causal effects of income, infant mortality, and education, 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 stylized facts. These derive from the application of Phillips and Sul's (2007) methods to the annual time series of fertility rates of 180 countries from 1950 to 2015. The nonlinear and semiparametric approach to transition modeling allows quite general time paths and heterogeneity across countries. In this setting, convergence means that, in the long run, fertility rates of countries in a cluster follow the same stochastic trend.

The identification of convergence groups is useful to pick out those countries that experienced a fertility transition after World War II. We concentrate on this group of countries to estimate the determinants of the decline of fertility in developing countries. According to demographers (e.g., Dyson and Murphy, 1985; 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 characterized by the clear distinction between pre- and posttransition regimes. For these reasons, we investigate the causes of fertility transitions from the point of view of the onset timing. We estimate the onset year of the fertility transition by applying the recent econometric methods proposed by Pierre Perron and his coauthors (Perron and Zhu, 2005; Perron and Yabu, 2009). These methods provide an estimation of the 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 noninferential procedures based on the rate of change of fertility (e.g., Coale and Watkins, 1986; Guinnane, Okun and Trussell, 1994; Reher, 2004).

We study the determinants of the onset of fertility transitions in 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 aim with this paper is to enhance understanding of the causes of demographic transitions. Accordingly, our investigation focuses on the main factors that have been debated in the literature: income, education, and child mortality (see Schultz, 1997; Guinnane, 2011; Galor, 2011; Galor, 2012). We also control for important factors often discussed in studies of the fertility decline in the developing world: the rate of urbanization, 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 behavior.

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. In this context, we use both traditional estimators based on external instruments and the approach recently proposed by Lewbel (2012), where identification relies on IV generated from exogenous regressors and heteroskedastic random errors. In both cases, we test the validity of the IV using the recent methods that extend weak-instrument inference to models with multiple endogenous regressors (Stock and Yogo, 2005; Angrist and Pischke, 2009; Sanderson and Windmeijer, 2016). In the traditional IV approach to identification, we instrument percapita 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, 2015). 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. This variable captures both important drug and chemical innovations, particularly efficient against epidemic diseases, and innovations in the international institutions for health promotion around the world. The instrument for the years of primary 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 econometric analysis of a 180-country panel for the period from 1950 to 2015 reveals two main results. The investigation of the time paths of total fertility rates (TFR) highlights some important features of global fertility transitions. The most immediate result concerns 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 decline. 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 characterized 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 then identify countries where a fertility transition occurred after 1955 and investigate the causes that triggered the transition. This sample includes 75 countries with available data on the variables listed above as possible determinants of transitions. 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 9.7 months earlier than without such increase, while a 10% reduction in infant mortality implies the same year comes around 7.9 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, from a different point of view, confirm 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 Theoretical and Empirical Literature

2.1 Theory

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 (Cleland, 2001). From this perspective, it must be noted that, in a demographic transition, we observe not just a decline in total fertility, but also a reduction of fertility net of mortality of children in the early months or years of life. 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; Azarnert, 2006). An analysis of the effect of child mortality on fertility depends on the assumptions made in the literature on some basic features of the household decision problem.

When theory admits uncertain child survival, a precautionary demand for children can be the outcome of the parents' decision problem. 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.¹ 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). Becker distinguishes quantity and quality of children in the parents' decision problem. 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.²

The influential Unified Growth Theory (UGT) argues the demographic transition is

¹Doepke (2005) shows that a positive relationship between child mortality and fertility is subject to the condition that risk aversion with respect to surviving children is stronger than risk aversion with respect to consumption.

 $^{^{2}}$ Galor (2012) assesses this approach.

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 takeoff. 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 labor than to physical labor, and women have a comparative advantage in mental labor. The decline in the gender wage gap causes the decline of fertility. The rise of women's relative wages is a stylized fact of economic growth since the nineteenth century.³

2.2 Empirics

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 regularities, and find remarkable variation across nations of both the onset and the pace of fertility decline (e.g., Dyson and Murphy, 1985; 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. The present paper investigates fertility convergence in the world using Phillips and Sul's (2007) panel time-series methods, which allow for heterogeneity across individuals and over time. We find that the world distribution of TFR is characterized by three

³Recently, an interesting research line focuses on the role of women empowerment in the demographic transition (e.g., de la Croix and Vander Donckt, 2010; Diebolt and Perrin, 2013).

convergence groups in the 2000s.

The largest part of the literature on the causes of fertility decline focuses on indicators such as TFR and crude birth rate (see the surveys of Schultz, 1997; Galor, 2012). Recently, two studies take this paper's point of view, estimating models of the timing of the onset of the fertility transition. Spolaore and Wacziarg (2014) test the hypothesis that the timing of the fertility decline in European regions has been determined by a process of diffusion of culture and social norms from the first country that adopted a new model of reproduction, France, to its imitators. They estimate regression models for the marital fertility transition date (Coale and Watkins, 1986) and a binary indicator of the transition in each 20-year interval from 1841 to 1941. The results support the main hypothesis advanced in the paper. Basso and Cuberes (2017) study the determinants of the onset year of the fertility transition using the genetic distance from the world technological frontier (UK and USA) as a proxy of the degree of international technology adoption. They find the genetic distance contributes to the explanation of the timing of fertility declines.

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, DDT⁴) and by the adoption of a new international health policy with the establishment of the World

⁴Dichlorodiphenyl trichloroethylene.

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. Accemoglu and Johnson find a positive relation between life expectancy and births.

Investigating the same research problem, Lorentzen et al. (2008) instrument both adult and infant mortality with the Malaria Ecology Index (Kiszewski et al., 2004).⁵ This index has been calculated on many countries, but is time invariant.⁶ Lorentzen et al. (2008) use this variable and other geographic and climate indicators to instrument adult mortality, infant mortality, secondary school enrollment, in estimates of the crosscountry 1960–2000 average of the rate of fertility. Both mortality indicators display positive coefficients while the effect of school enrollment is negative. 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.

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 fertility 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 enrollment rates are instrumented with the ratios of the numbers of telephones, radios, and newspapers to adults in the population. Scotese Lehr finds secondary enrollment rates have a negative effect on net fertility.

 $^{^{5}}$ The malaria ecology index has been derived with the aim of measuring the basic potential transmission of the disease at regional level in the world. This measure should be plausibly exogenous to economic conditions and public health interventions.

⁶The data of each country are constructed using the mean monthly temperature from 1901 to 1990.

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. 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.

Herzer et al. (2012) propose an innovative approach based on Granger causality and panel cointegration. The authors investigate the long-run equilibrium relation linking the crude birth rate to per-capita GDP and the crude mortality rate in 20 countries from 1900 to 1999. They find fertility has a negative relationship with per-capita income and a positive relationship with mortality. Tests of the weak exogeneity of the three variables lead to a rejection of the hypothesis, which means they probably are endogenous. Impulseresponse analysis illustrates the relations among the variables. Cointegration methods have the significant advantage that the estimated long-run equation remains valid even if other relevant variables are excluded. The application of this methodology to panel data from 119 countries over the period 1950–1999 confirms the main results of the paper.

3 Modeling 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:

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

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

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

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

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 modeled by the regression equation

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

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 lefthand side of the regression equation (5) diverges to $-\infty$. We perform a one-sided test of the null hypothesis of convergence ($\gamma \ge 0$)) 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. 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 and Prescott's (1997) 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:

$$\log\left(\frac{H_1}{H_t}\right) - 2\log\left[\log\left(t+1\right)\right] = \underset{(2.36)}{0.609} - 1.092\log\left(t\right),$$

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 = -10.24. 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 (Wilson, 2011). The alternative to global convergence could be the case where every country has a distinct long-term TFR value, or, more realistically, the case of convergence within groups. The investigation of such an alternative configuration of fertility dynamics can be pursued by applying Phillips and Sul's (2007) clustering algorithm to the TFR data.

The procedure starts by ordering the countries according to the last five years' average TFR. The first country in the list is Niger with the highest fertility rate (7.37), while the last is Singapore (TFR=1.23). In the second step, a core group of countries is formed. This step requires the selection of the first k countries with a log t test that does not reject the null (i.e., at the 5% level, $t_k > -1.65$). Within this group, the core is formed by those countries that display the highest value of t_k .⁷ The core is formed by Niger, Somalia and

 $k^* = \arg\max_k \{t_k\}$ subject to $\min\{t_k\} > -1.65$.

⁷The core group size k^* is given by the criterion:

the Democratic Republic of Congo $(t_k = 1.13)$. The third step of the algorithm is the formation of the first convergence group. Each country is added to the core group to run the log t test that provides a statistic denoted \hat{t} . Those countries with $\hat{t} > c$, where c is a critical value, enter the convergence group. The critical value we choose, c = 0, is quite conservative, as Phillips and Sul (2007) recommend.⁸ This process leads to the formation of a first cluster that includes the core countries, Chad and Timor-Leste. Repeating the clustering procedure again, we find Mali does not satisfy the condition $\min\{t_k\} > -1.65$ that allows the formation of a core group among the remaining countries. However, the log t test says that the first group and Mali can be merged: $t_k = 3.03$. Repeating the procedure again we find another group of eight countries that can be merged with the first in a group of 14 countries converging to a high fertility rate (Group H). The log t test for the whole Group H provides a point estimate of $\hat{\gamma} = -0.091$, and a test statistic of $\hat{t} = -0.57$. 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.

The application of the log t test to the remaining countries does not support the convergence hypothesis since $\hat{t} = -7.74$. Hence, we repeat the first three steps of the algorithm on data for those countries. In this way, we identify another convergence group of 25 countries. For this group, the log t test does not reject the null ($\hat{t} = -0.42$), although the evidence for convergence is weak. 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

⁸Thsi conservative testing approach could lead to the identification of more groups than the true number. To avoid this outcome of the clustering procedure, Phillips and Sul (2009) suggest the use of the log t test to assess the hypothesis that two groups can be merged in a larger group. Here, we follow their suggestion.

 $(\hat{t} = -0.24)$, although its negative value indicates a very 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.

Hence, with the application of the clustering algorithm we have found three convergence groups. 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 coefficients $\hat{\gamma} = -1.32$ and $\hat{t} = -20.94$.

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. 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 Cluster H-L which also includes many other countries of Latin America, Asia, Africa, and Oceania. Table 2 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 provides several insights into the process that drove the dramatic change in reproductive behavior in many developing countries after World War II. Indeed, Figure 1 clearly shows the main features of the dynamic path of fertility and highlights sharp heterogeneity across countries both in the extent of the transition and in its timing. In this respect, we can distinguish three fundamental patterns in the world since the 1950s: a small cluster of countries where fertility remained steadily high, a significant number of countries with initially low and then decreasing TFR, and a large group of countries undertaking a dramatic reduction of fertility starting from a number of children per woman close to six and ending with a value close to two. The last phenomenon 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

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 .

$$y_t = \mu + \beta_1 t + \beta_b B_t + u_t, \qquad t = 1, ..., T$$
 (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 :

$$B_{t} = \begin{cases} 0 & if \quad t \leq T_{1}, \\ t - T_{1} & if \quad t > T_{1}. \end{cases}$$

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 . Since Herzer et al. (2012) provide econometric evidence of a stochastic trend in fertility time series, the Perron and Zhu's methods can be correctly applied in the present context. 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 at least 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. The plot of the TFR time series suggests that fertility transitions in many countries can be represented by a smooth trend with an important break. Nevertheless, we find in the data some deviations from this stylized picture. Some countries have reached a very low TFR some years before 2015 and display constant fertility or a slightly increasing trend in the years after the transition. In some Sub-Saharan countries fertility rates rose again during the 1980s and 1990s. This phenomenon could be the consequence of the higher mortality caused by the AIDS epidemic. In both cases, the decline of fertility can also be followed by a weaker trend with the opposite sign. More generally, occasional deviations from the main trends are not rare. We do not consider these minor forms of change in the trends and focus on the main distinction between the periods before and after the onset of the demographic transition. Actually, we have used Phillips and Sul's (2007) methods to select those countries where this choice is clearly justified. To ensure we estimate the year that correctly indicates the onset of a significant fertility decline, we limit the search for T_1 to those models (6) where the estimate of β_b is negative. 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 3 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} . Since our dependent variable is the logarithm of TFR, β_1 and β_2 can be interpreted as the average rate of change of TFR during the period respectively before and after the break year T_1 . Countries in Table 3 were ordered by group (I and H-L) and by TFR (high to low) in the last five years of the interval 1950–2015. This ordering highlights the strength of the fertility transition across countries. For each country, the plot of the time series of TFR is shown in Appendix. In each graph, a vertical line represents the estimate of the onset year. Table 3 does not show the results of Aruba and Singapore because their TFR always decreases between 1956 and 2009.

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 progress. 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.

 $^{^9\}mathrm{Reher}$ (2004) estimates the onset of fertility decline in China in 1970, using quinquennial data on births.

5.1 Econometric Model and Data

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

$$onset_{it} = \alpha \log(GDP_{it}) + \beta \log(infant \ mortality_{it}) + \gamma \log(schoolw_15 - 39_{it}) + (7)$$
$$+ \mathbf{x}_{it}\boldsymbol{\delta} + \mu_t + c_i + u_{it}, \quad i = 1, ..., N, \quad t = 1, ..., T,$$

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, **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. 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 3. 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 Maddison (2001) updated by 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,¹⁰ 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. 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., de la Croix and Vander Donckt, 2010; Diebolt and Perrin, 2013) 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. Summary statistics are in Table 4.

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. In this regard, equation (7) can be interpreted as part of a structural model. 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 methodol-

¹⁰to save the life of the woman; to preserve her physical health; to preserve her mental health; consequent on rape or incest; fetal impairment; economic or social reasons; and available on request (Bloom et al. 2009, p. 87).

ogy. 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. More, 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 purges 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. Accemoglu and Johnson (2007) use predicted mortality as an instrument for life expectancy in two-stage least squares (2SLS) estimates of the total number of births and find a significant positive effect. Similarly, Hansen and Lønstrup (2015) extend Accemoglu and Johnson's (2007) data with data in 1900, and find a positive effect of life expectancy on population size. 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 derived from the international oil price (*oil shock*), and they construct "a country-specific oil price shock variable as the change in the log of the international oil price weighted with countries' sample average netexport shares of oil in GDP" (Brueckner and Schwandt, 2015, p. 1654). This variable 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.¹² Barro and Lee (2015) follow this approach to investigate the determinants of fertility with a special focus on the role of education. 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 mothers' education in equation (7)

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

 $^{^{12}}$ A recent example is Barro (2015), while two recent surveys of the literature on education and economic growth are in Glewwe et al. (2014) and Barro and Lee (2015).

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 heteroskedasticity and clustered at the country level. Estimates of model (7), shown in Table 5, 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 highly 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. A significant result in our estimates – which will be confirmed in the following analysis – is the prominent role of urbanization. Indeed, since the main four regressors enter the estimates after a logarithmic transformation, their coefficient is the change in the probability of the onset of the fertility transition that is associated with a 1% increase in the explanatory variable. These figures can be compared. The urbanization rate stands as one of the most important correlates of the onset of a fertility decline.

The preceding estimation results show important correlations between the onset probability of the fertility transition and some of the variables suggested in the literature as this phenomenon's main determinants. The coefficients' signs are strictly consistent with those implied by the theory. However, OLS regressions do not enable us to establish causal relationships. We tackle this issue in the following sections.

5.2.1 IV Estimation: One Endogenous Regressor

In this section we present the results of the estimation of IV models where only one of the three endogenous explanatory variables is considered in turn. This is 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). In this paper, the approach is justified by the use of IV whose validity and relevance have been established in similar econometric contexts (see the references above), and by the advantages it offers. Indeed, IV estimators are consistent and asymptotically normal. Hence, we get coefficient estimates with a significant statistical content that adds to that obtained by estimating models with the full set of endogenous regressors. Furthermore, estimates relying on different instruments could inform us about their robustness.

Here, we present the results of panel, fixed-effects, 2SLS estimates¹³ of models of the dependent variable onset where the set of regressors includes in turn *GDP*, *infant mortality*, *schoolw_15–39*. Regressions also include the time-varying variables *urbanization* and *abortion*. All the right-hand-side variables enter the estimates as log transformations. Table 6 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. Angrist and Pischke (2009)

 $^{^{13}}$ We used the unofficial STATA command ivreg2 by Baum, et al. (2010). See also Baum, et al. (2007b).

argue that the 2SLS estimator of exactly identified models is approximately unbiased, even when instruments are weak. 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 the exactly identified models, while the estimates are negative and significant in the last three overidentified equations. The variable *urbanization* displays positive and significant coefficient estimates in all the specifications. In the first stage regression results of Table 6, the IV coefficients show significant estimates of the expected sign, and the Kleibergen–Paap Wald F-statistic (Kleibergen and Paap, 2006) 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 7 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. Their absolute value is close to 0.6, meaning that if the rate of infant mortality decreases by 10%, the probability of a fertility transition is predicted to increase by 6%, which is a large effect. The first-stage estimates in these 2SLS regressions highlight the relevance of the chosen instruments, with t-statistics 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.

The last application of the one-endogenous-regressor approach refers to mothers' educational attainment. Table 8 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 log *urbanization* are strongly significant (at the 1% level) 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. Here, we implement an alternative strategy: 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 GMM (Baum et al., 2007; Stock and Yogo, 2005). Table 9 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 13 variables: *world income*, plus two variables obtained from the interactions between *world income* and dummy variables 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 dummy variables 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 set of interactive instruments capture the distinctive impact of the international epidemiological transition (Acemoglu and Johnson, 2007) 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 9, 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 -0.31 to -0.42, 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.

The first-stage results in Table 10 show the estimates of the coefficients of the IV. 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. 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) show 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. 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 of the null: The endogenous regressors in equation (7) are irrelevant. This test is robust to weak instruments.¹⁴ In all the second-stage 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.

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-endogenousregressor 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 traditional exclusion restrictions. Indeed, Lewbel's (2012) method is based on IV generated from variables exogenous to the model. Let us consider equation (7) in

 $^{^{14}\}mathrm{Baum}$ et al., (2007) discuss this test.

the setting of a model for cross-country data:

$$transition_{i} = \alpha' \log(GDP_{i}) + \beta' \log(infant \ mortality_{i}) + \gamma' \log(schoolw_15 - 39_{i}) +$$

$$(8)$$

$$+\mathbf{x}_{it}\boldsymbol{\delta}' + c + \epsilon_i, \quad i = 1, ..., N,$$

Identification rests on \mathbf{z} , a vector of exogenous variables that may be equal to \mathbf{x} or may include other exogenous variables.¹⁵ 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 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.¹⁶ 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 application of Lewbel's (2012) 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 estimates, the regressors are averages across

¹⁵Lewbel (2012) assumes that variables not in \mathbf{x} may enter \mathbf{z} . The variables of \mathbf{z} can be discrete or continuous.

¹⁶The method of Lewbel (2012) has been applied in several papers including, more recently, Emran and Hou (2013); Chowdhury et al. (2014); Millimet and Roy (2016).

time of the variables in model (7). 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 8.0.

Table 12 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 the usual significance levels.

To compare the marginal effect of the logged variables with that of the untransformed variables, Table 12 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 effect of *schoolw_15-39* is 0.15: if the educational attainment of women increases by 10% the year of the onset comes 9.7 months earlier, while the corresponding coefficients for infant mortality are -0.122 and 7.9 months, respectively.¹⁷

 $^{^{17}\}mathrm{We}$ calculate the effect of a change in an explanatory variable x on the onset year T_b considering that:

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 (Sanderson and Windmeijer, 2016) takes values from 23 to 108. 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 12 presents the results of the application of the Ramsey RESET test proposed by Pesaran and Taylor (1999) for IV estimators.¹⁸ In all regressions, the null hypothesis that there are no neglected nonlinearities can be comfortably accepted.

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, both of which occurred in the period for much of the world population. Regression analysis also produces evidence consistent with the view that children are

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

¹⁸We used the unofficial STATA command *ivreset* by Mark E. Schaffer: https://ideas.repec.org/c/boc/bocode/s455701.html

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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Appendix

Phillips and Sul's (2007) Approach

Recently, panel time-series econometrics has seen the development of factor models where unobserved common factors and idiosyncratic components can be distinguished. The case of a single-factor model for the variable X_{it} for individual *i* at time *t* is:

$$X_{it} = \delta_i \mu_t + \epsilon_{it},$$

where μ_t is the common factor, δ_i is a factor loading specific of individual *i*, and ϵ_{it} is a random error. This model assumes that the distance between the common factor and the systematic part of X_{it} is specific of the individual *i* but is constant over time. Phillips and Sul (2007) extend this model with the following factor representation which allows for a time varying δ_{it} that absorbs ϵ_{it} :

$$X_{it} = \delta_{it}\mu_t,\tag{A1}$$

where δ_{it} represents the transitional path of the variable to the common trend component. This path is time varying and specific of unit *i*. Phillips and Sul (2007) model the transitional components in a semiparametric form:

$$\delta_{it} = \delta_i + \sigma_i \xi_{it} L(t)^{-1} t^{-\alpha}, \ \sigma_i > 0, \tag{A2}$$

where δ_i is fixed, ξ_{it} is iid(0, 1) across i and weakly dependent over t, and L(t) is a slowly varying function¹⁹, such that $L(t) \to \infty$ as $t \to \infty$. In model (A2), δ_{it} converges to δ_i for all $\alpha \ge 0$. Interestingly, the model accommodates heterogeneous transition paths both across individuals and over time²⁰.

This nonlinear, time varying factor model provides the basis for the proposal of a convergence test and a clustering procedure (Phillips and Sul, 2007). Convergence among the series X_{it} is defined as the long run equilibrium of their ratios:

¹⁹A slowly varying function F(t) satisfies the condition: $F(at)/F(t) \to 1$ as $t \to \infty$ for all a > 0.

²⁰Phillips and Sul (2007) show how their framework is even more general because it allows the parameter α and the function L(t) to be specific of individual *i*.

$$\lim_{k \to \infty} \frac{X_{it+k}}{X_{jt+k}} = 1 \text{ for all } i \text{ and } j.$$
(A3)

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 *logt* test of convergence is defined in terms of the relative transition coefficients:

$$h_{it} = \frac{X_{it}}{\frac{1}{N} \sum_{i=1}^{N} X_{it}} = \frac{\delta_{it}}{\frac{1}{N} \sum_{i=1}^{N} \delta_{it}},$$
(A4)

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

$$\lim_{t \to \infty} \frac{1}{N} \sum_{i=1}^{N} (h_{it} - 1)^2 = 0.$$
 (A5)

Equations (A1) and (A2) with further assumptions detailed in Phillips and Sul (2007) allow the proposal of a test of the null hypothesis of convergence:

$$H_0: \delta_i = \delta$$
 and $\alpha \ge 0$,

against the alternative:

$$H_A: \delta_i \neq \delta$$
 for all *i* or $\alpha < 0$.

It can be noted that the alternative hypothesis is quite general because it admits the possibility of club convergence or local convergence to multiple long-run equilibria. The test is based on the cross-sectional variance:

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

The dynamics of the ratio $\frac{H_1}{H_t}$ over time are explained by the regression equation

$$\log\left(\frac{H_1}{H_t}\right) - 2\log L\left(t\right) = a + \gamma\log\left(t\right) + \hat{u}_{t,} \tag{A6}$$

for t = [rT], [rT] + 1,,T with [rT] the integer part of rT and r the fraction of the time-series observations discarded in the regression. L(t) is a slowly varying function as $L(t) = \log(t+1)$ suggested by Phillips and Sul (2007). They show that the estimate of the coefficient γ converges in probability to $2\hat{\alpha}$, where $\hat{\alpha}$ is the estimate of α under the null. Convergence implies 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 dependent variable of the regression equation (A6) diverges to $-\infty$. A one-sided test of the null hypothesis $\alpha \geq 0$ is carried out using the t-statistic calculated using the estimate of γ and a HAC standard error. This statistic is asymptotically distributed as a standard normal.

Significantly, the log t test is consistent even when the alternative hypothesis maintains that the transition coefficients δ_{it} converge but to different δ_i . Hence, club convergence is admitted as a form of divergence. Phillips and Sul (2007) propose a procedure to identify long-run clusters of convergence. The algorithm starts by ordering all units in the panel according to the value of the last observation available 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.

Perron and Yabu's (2009) Test for Structural Change in the Trend Function

Perron and Yabu (2009) propose a test of the null hypothesis of one break in the slope of the trend function (6). Under the null $\beta_b = 0$, meaning the trend function (6) does not change in the period under consideration. The test procedure can be applied to models as (6) under the same general assumptions on the dynamic properties of the noise component that admit either stationarity or a unit root in u_t . In particular, equation (6) can be expressed in matrix notation as

$$y_t = \mathbf{x}_t' \Psi + u_t,$$

where $\mathbf{x}_t = (1, t, B_t)'$ and $\Psi = (\mu, \beta_1, \beta_b)$. The Wald test W_{RQF} (Wald robust quasi-FGLS

test) derives from the estimation of the quasidifference equation,

$$(1 - \alpha L) y_t = (1 - \alpha L) \mathbf{x}'_t \Psi + (1 - \alpha L) u_t,$$
(B1)

where α is the sum of the autoregressive coefficients of the model for the random component u_t with unknown order k. The estimation of equation (B1) requires an estimate of α that can be obtained from the OLS regression:

$$\widehat{u}_t = \alpha \widehat{u}_{t-1} + \sum_{i=1}^k \zeta_i \Delta \widehat{u}_{t-i} + e_t, \tag{B2}$$

where \hat{u}_t is the residual from the OLS estimation of model (B2). Perron and Yabu (2009) basic framework can be applied under the assumption $|\alpha| \leq 1$, with two corrections to the estimate of α . The first, provides the super-efficient estimate,

$$\widehat{\alpha}_{S} = \begin{cases} \widehat{\alpha} & if \quad T^{\delta} \left| \widehat{\alpha} - 1 \right| > d \\ 1 & if \quad T^{\delta} \left| \widehat{\alpha} - 1 \right| \le d \end{cases}$$

for $\delta \in (0, 1)$ and d > 0. Perron and Yabu's (2009) Theorem 1 asserts that the Wald test obtained using $\hat{\alpha}_S$ has the same limit distribution, χ^2 , in both cases: $|\alpha| < 1$ and $\alpha = 1$. The other correction reduces the finite-sample bias in the OLS estimate of α when the coefficient is close to one. Here, Perron and Yabu adopt the method proposed by Roy and Fuller (2001).

Group H	Group I	Group H-L		Group L-L
Niger	Tanzania	Samoa	Cabo Verde	Gabon
Somalia	Afghanistan	Ghana	Suriname	Israel
Congo, Dem. Rep.	Benin	Kenya	India	Argentina
Mali	Zambia	Solomon Islands	Indonesia	New Caledonia
Chad	Cote d'Ivoire	Zimbabwe	Guam	Jamaica
Burundi	Guinea	Papua New Guinea	Venezuela, RB	Uruguay
Angola	Central African Republic	Tonga	Libva	New Zealand
Uganda	Senegal	Pakistan	Uzbekistan	France
Timor-I este	Cameroon	Iordan	Nenal	Georgia
Nigeria	Malawi	Namihia	Nicaragua	A zerbaijan
Burking Eaco	Cuinea Bissau	Taiileistan	Muanmar	Iroland
Cambia The	Mouritonio	Vapuatu	Marrian	Iceland
Gambia, The	Maulitallia Conso Dor	Valluatu	Turicio	Kerrer Dem Beenle's Den
Mozambique	Congo, Rep.	Egypt, Arab Rep.	Tunisia	Korea, Dem. People's Rep.
Equatorial Guinea	Sierra Leone	Micronesia, Fed. Sts.	Bangladesh	Sweden
	Liberia	Swaziland	Bhutan	Australia
	Sudan	Lesotho	Maldives	United States
	Togo	Guatemala	Grenada	United Kingdom
	Sao Tome and Principe	Haiti	El Salvador	Chile
	Comoros	Syrian Arab Republic	Sri Lanka	Bahamas, The
	Ethiopia	Djibouti	Turkey	Norway
	Iraq	Philippines	Bahrain	Barbados
	Madagascar	Bolivia	Malaysia	Belgium
	Eritrea	Turkmenistan	French Polynesia	Finland
	Vemen Den	Algeria	Kuwait	Netherlands
	Pwanda		Oatar	Montenagro
	Kwalida	LaorDK	Qalai	Dermart
		Bolswana	Vietnam	Ortha
		Oman		Cuba
		Mongolia	Brunei Darussalam	Russian Federation
		Saudi Arabia	Costa Rica	Armenia
		Cambodia	United Arab Emirates	Belarus
		Honduras	Aruba	Canada
		Belize	Trinidad and Tobago	Lithuania
		Fiji	Brazil	Slovenia
		Paraguay	Iran, Islamic Rep.	Estonia
		Guyana	Albania	Switzerland
		Panama	Lebanon	Luxembourg
		Ecuador	China	Latvia
		Morocco	Thailand	Bulgaria
		South Africa	Mauritius	Likraine
		Dominican Penublic	Korea Den	Macedonia EVP
		Port	Singapora	Puerto Pico
		Telu	Singapore	Pomonio
				Komania Geogle Denselation
				Croatia
				Channel Islands
				Austria
				Germany
				Japan
				Malta
				Italy
				Cyprus
				Slovak Republic
				Hungary
				Greece
				Bosnia and Herzegovina
				Doland
				rolaliu
				0
				Spain
				Spain Moldova

Table 1. Countries in convergence groups

	Countries	Min	Min	Max	Max	Mean	Mean
		1950	2015	1950	2015	1950	2015
Group H	14	5.23	4.77	7.34	7.29	6.53	5.82
Group I	25	5.40	3.97	8.11	5.08	6.57	4.63
Group L	141	1.87	1.23	7.68	4.04	5.09	2.18
of which:							
Group H-L	82	4.89	1.24	7.68	4.04	6.42	2.51
Group L-L	59	1.87	1.23	5.52	3.85	3.25	1.72
World	180	1.87	1.23	8.11	7.29	5.41	2.80

Table 2. Indicators of fertility in convergence groups

Figure 1. Total Fertility Rate in Convergence Groups, 1950-2015



Table 3. Estimate of the onset year of fertility transition

Country	Onset	W _{RQF}	β_1	β_2	Country	Onset	W _{RQF}	β_1	β ₂
	year					year	•		
Group I									
Tanzania	1977	47.781	0.000	-0.007	Sierra Leone	1995	45.709	0.003	-0.019
Afghanistan	2002	31747.139	0.000	-0.033	Liberia	1984	53.235	0.004	-0.013
Benin	1982	57.131	0.007	-0.011	Sudan	1979	52.096	0.002	-0.011
Zambia	1973	59.892	0.005	-0.009	Togo	1978	61.892	0.006	-0.013
Cote d'Ivoire	1974	58.709	0.003	-0.012	Sao Tome and Principe	1978	52.433	0.002	-0.010
Guinea	1992	41.873	0.003	-0.013	Comoros	1979	2.185	0.006	-0.015
Central African Republic	1980	47.328	0.003	-0.005	Ethiopia	1995	46.328	0.001	-0.026
Senegal	1979	54.238	0.003	-0.013	Iraq	1974	3.827	0.002	-0.013
Cameroon	1984	56.540	0.007	-0.011	Madagascar	1978	50.497	-0.001	-0.013
Malawi	1983	56.352	0.004	-0.014	Eritrea	1989	44.055	-0.002	-0.017
Guinea-Bissau	1990	47.814	0.004	-0.014	Yemen, Rep.	1989	72.098	0.006	-0.031
Mauritania	1971	2.702	0.004	-0.008	Rwanda	1983	63.864	0.002	-0.023
Congo, Kep.	1972	59.449	0.006	-0.007					
Country	Onset year	W _{RQF}	β_1	β2	Country	Onset year	W _{RQF}	β_1	β_2
Group H-L									
Samoa	1963	64.851	0.001	-0.013	Peru	1965	71.110	0.000	-0.023
Ghana	1973	63.050	0.004	-0.014	Cabo Verde	1982	70.863	-0.001	-0.032
Kenya	1973	66.157	0.004	-0.017	Suriname	1959	74.293	-0.001	-0.020
Solomon Islands	1977	64.152	0.005	-0.016	India	1973	62.266	-0.003	-0.020
Zimbabwe	1972	64.509	0.006	-0.019	Indonesia	1966	70.268	0.002	-0.021
Papua New Guinea	1970	58.536	0.000	-0.011	Guam	1956	233.205	0.005	-0.016
Tonga	1957	1917.353	-0.001	-0.012	Venezuela, RB	1960	75.358	0.001	-0.020
Pakistan	1985	55.042	0.000	-0.022	Libya	1975	79.764	0.007	-0.037
Jordan	1973	64.386	0.004	-0.023	Uzbekistan	1968	4.919	0.016	-0.026
Namibia	1977	66.789	0.005	-0.020	Nepal	1992	70.418	-0.003	-0.040
Tajikistan	1967	91.665	0.021	-0.018	Nicaragua	1978	67.203	-0.005	-0.032
Vanuatu	1997	24.621	-0.012	-0.020	Myanmar	1969	71.679	0.001	-0.023
Egypt, Arab Rep.	1972	49.241	-0.004	-0.019	Mexico	1968	72.943	0.000	-0.027
Micronesia, Fed. Sts.	1975	58.154	-0.002	-0.019	Tunisia	1968	68.416	0.004	-0.031
Swaziland	1981	62.774	0.000	-0.023	Bangladesh	1976	85.205	0.004	-0.033
Lesotho	1982	54.080	-0.001	-0.019	Bhutan	1986	81.635	-0.001	-0.042
Guatemala	1988	45.826	-0.005	-0.025	Maldives	1983	86.508	0.005	-0.046
Haiti Gurian Anala David II	1988	54.466	-0.002	-0.025	Grenada	1957	4050.247	0.036	-0.021
Syrian Arab Republic	1977	69.805	0.001	-0.026	El Salvador	1970	72.902	0.001	-0.026
	1987	66.272	0.001	-0.030	Sri Lanka	1956	1282.938	-0.007	-0.019
Philippines	1961	66.985	-0.004	-0.017	l urkey	1966	62.897	-0.008	-0.024
BOIIVIa	1982	48.874	-0.007	-0.020	Banrain	1966	76.988	0.002	-0.027
i urkmenistan	1964	91.099	0.026	-0.022	Ivialaysia Frenck Deliversit	1960	/6.0/4	0.001	-0.021
	1000	57.451	0.003	-0.031	French Polynesia	1969	59.170	-0.009	-0.022
	1076	/5.959 65 410	0.002	-0.035	Nuwait	1075	4.200	0.001	-0.030
Oman	100E	75 216	0.000	-0.025	Vietnam	1050	12.204 QG 211	0.002	-0.032
Mongolia	1000	/5.210 60 770	0.004	-0.043	Colombia	1060	80.241 77 375	0.014	-0.033
Saudi Arabia	1083	77 654	0.023	-0.033	Rrunei Darussalam	1963	68 051	-0.002	-0.020 -0.028
Cambodia	1000	12.004 57 702	-0.000	-0.033		1057	1203 E00	0.002	-0.028
Honduras	1070	57.703 61.605	-0.005	-0.032	Linited Arab Emirator	1021	67 157	-0.020	-0.023
Relize	1979	65 000	-0.004	-0.027	Trinidad and Tohago	1956	3158 447	-0.007	-0.039
Fiii	1956	25 760	-0 012	-0.024	Brazil	1962	72 007	-0.002	-0 022
'יני Paraguay	1950	<u>44</u> 78 <u>4</u>	-0.013	-0.010	Iran, Islamic Ren	1983	73.992	-0.001	-0.020
Guvana	1056	44.704 456524 QAA	0.010	-0.025	Albania	1965	A1532 AAD	0.004	-0.049
Panama	1950	320 24.500	0.019	-0.018	Lehanon	1958	63 U88	-0.013	-0.027 -0.028
Fruador	1958	69 727	-0 001	-0.017	China	1963	03.000	-0.004	-0.028
Morocco	1060	60 7/2	0.001	-0.021	Thailand	1062	7/ /25	0.000	-0.031
South Africa	1060	62 565	-0.003	-0.020	Mauritius	1902	11 876	-0.002	-0.032
Dominican Republic	1958	1075 737	0.001	-0.021	Korea, Ren	1956	15 084	0.034	-0.034
Bommean Republic	100	1073.737	0.001	0.022	Norca, nep.	1001	10.004	0.004	0.004

Table 4. Summary statistics for the variables of the transitioning countries

Variable	Observations	Countries	Mean	S.D.	Min	Max
Onset	675	75	0.636	0.482	0	1
TFR	675	75	5.596	1.544	1.501	8.800
GDP	675	75	3.261	4.555	0.381	37.112
Infant mortality	675	75	92.406	50.346	4.080	319.594
Schoolw_15-39	675	75	2.552	1.642	0.019	7.082
Urbanization	675	75	36.577	22.363	2.049	100
Catholic	146	73	27.170	34.735	0	96.600
Muslim	146	73	30.155	38.544	0	100
Women vote	73	73	1954.205	15.947	1920	2006
Abortion Index	519	71	2.399	2.042	0	7

Correlation Matrix	Obs.	Onset	Log(GDP)	Log(inf.mortality)	Log(schoolw_15-	Log(urbaniz.)	Log(abortion)
					39)		
Onset	675	1					
Log(GDP)	675	0.22	1				
Log(inf. mortality)	675	-0.50	-0.64	1			
Log(schoolw_15-39)	675	0.60	0.36	-0.65	1		
Log(urbanization)	675	0.38	0.70	-0.55	0.41	1	
Log(abortion)	519	0.16	0.07	-0.21	0.15	0.08	1

Table 5. Determinants of fertility transition onsets. Dependent variable: *onset*. Panel fixed and random effects

	(1) (2)		(3)	(4)	(5)	(6)
-	Fixed	Effects		Random	Effects	
Log(GDP)	-0.202****	-0.161**	-0.196***	-0.169***	-0.163***	-0.137***
	(-4.077)	(-2.314)	(-5.234)	(-4.010)	(-3.909)	(-2.864)
Log(infant mortality)	-0.141***	-0.007	-0.153**	-0.117*	-0.117*	0.004
	(-2.677)	(-0.098)	(-2.268)	(-1.724)	(-1.683)	(0.047)
Log(schoolw_15-39)	0.244***	0.334***	0.230***	0.258***	0.256***	0.334***
	(4.950)	(7.187)	(5.349)	(5.261)	(5.170)	(6.561)
Log(urbanization)	0.435***	0.438***	0.264***	0.247***	0.236***	0.241***
	(10.068)	(8.491)	(6.094)	(4.768)	(4.594)	(4.009)
catholic				-0.001	-0.001	-0.001
				(-0.774)	(-0.750)	(-0.600)
muslim				0.001	0.001	0.001
				(0.886)	(1.147)	(1.350)
women vote					-0.001	0.000
					(-1.099)	(0.078)
Log(abortion)		0.029				0.017
		(0.596)				(0.491)
Time dummies	YES	YES	YES	YES	YES	YES
	-					-
R^2	0.562	0.584	0.466	0.468	0.473	0.445
Observations	675	519	675	657	657	519
Countries	75	71	75	73	73	71

All regressions include time dummies. Standard errors are robust to heteroskedasticity and clustered at the country level. *t* statistics in parentheses. Columns 1-2 panel within estimates, columns 3-6 random effects estimates. p < 0.10, p < 0.05, p < 0.05, p < 0.01.

Table 6. Determinants of the onset of fertility transitions. Panel IV within estimates. Dependent variable: *onset*. Endogenous regressor: GDP per capita. Instruments: *oil shock*; *world income*.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Log(GDP)	0.085 (0.512)	-0.053 (-0.306)	0.115 (1.035)	0.432 (1.502)	0.106 (0.668)	0.063 (0.403)	-0.545*** (-3.846)	-0.383** (-2.237)	-0.357** (-2.278)
Log(urbanization)		0.613 ^{***} (5.459)	0.591 ^{***} (4.784)		0.643 ^{***} (5.087)	0.661*** (4.912)		0.547 ^{***} (4.221)	0.532*** (3.981)
Log(abortion)			-0.050 (-1.348)			0.017 (0.363)			-0.010 (-0.332)
First stage									
oil shock	2.642 ^{***} (4.159)	2.661*** (4.138)	2.369*** (2.677)				0.950 (1.521)	0.941 (1.516)	1.318 ^{**} (2.490)
world income				0.543*** (5.222)	0.452*** (8.165)	0.421*** (7.629)	0.403*** (6.277)	0.406 ^{***} (5.583)	0.407*** (5.322)
Hansen J J p-value K-P Wald test F Observations	17.300 426	17.124 426	7.165 384	27.273 531	66.669 531	58.195 427	0.577 0.447 20.805 376	0.195 0.659 17.001 376	0.308 0.579 19.176 346
Countries	64	64	61	59	59	57	54	54	52

Instruments: 1-3, oil shock; 4-6, world income, 7-9, oil shock and world income. All regressions include time dummies. Standard errors are robust to heteroskedasticity and clustered at the country level. K-P is the Kleibergen-Paap weak identification test statistic. *t* statistics in parentheses. * p < 0.10, ** p < 0.05, *** p < 0.01.

Table 7. Determinants of fertility transition onsets. Panel IV within estimates. Dependent variable: onset. Endogenous regressor: infant mortality. Instruments: predicted mortality; predicted mortality*dummy SSA

	(1)	(2)	(3)	(4)	(5)	(6)
Log(infant mortality)	-0.588*** (-5.221)	-0.561** (-2.206)	-0.642** (-2.383)	-0.587*** (-5.209)	-0.538** (-2.141)	-0.611** (-2.278)
Log(urbanization)		0.067 (0.151)	0.017 (0.039)		0.096 (0.223)	0.055 (0.127)
Log(abortion)			-0.064 (-0.977)			-0.061 (-0.928)
First stage						
predicted mortality	1.243*** (9.416)	0.806*** (4.293)	0.811 ^{***} (4.530)	1.242*** (8.987)	0.793 ^{***} (4.090)	0.794 ^{***} (4.292)
predicted mortality*dummy SSA				0.027 (0.139)	0.188 (1.083)	0.202 (1.350)
Hansen J J p-value				0.923 0.337	0.957 0.328	1.019 0.313
K-P Wald F-statistic	88.659	18.430	20.520	64.249	18.214	25.426
Countries	42	42	304 40	330 42	42	40

Instruments: 1-3, predicted mortality; 4-6, predicted mortality and predicted mortality interacted with dummy SSA countries. All regressions include time dummies.

Standard errors are robust to heteroskedasticity and clustered at the country level. K-P is the Kleibergen-Paap weak identification test statistic.

F-statistic refers to the Wald test on instruments in first-stage regressions. *t* statistics in parentheses. * *p* < 0.10, ** *p* < 0.05, *** *p* < 0.01.

Table 8. Determinants of fertility transition onsets. Panel IV within estimates.Endogenous regressor: Average years of schooling, female aged 15-39.

	(1)	(2)	(3)	(4)	(5)	(6)
Log(school_15-39)	0.470***	0.350***	0.353***	0.460***	0.332***	0.361***
	(11.342)	(6.494)	(6.585)	(9.964)	(5.322)	(6.691)
Log(urbanization)		0.313****	0.355****		0.332***	0.347***
		(4.716)	(4.887)		(4.534)	(4.765)
Log(abortion)			0.006			0.005
			(0.208)			(0.199)
First stage						
Log(school_40-64)	0.871***	0.724***	0.710****	0.844***	0.698***	0.702***
	(9.662)	(7.373)	(6.397)	(8.925)	(7.077)	(6.327)
Log(school 40-64)*dummySouthAsia				0.259**	0.251***	0.286***
				(2.460)	(3.124)	(3.234)
Hansen I				1.247	1.324	0.912
I p-value				0.264	0.250	0.340
K-P Wald F-statistic	93.361	54.354	40.916	189.969	114.025	44.901
Observations	674	674	518	674	674	518
Countries	75	75	71	75	75	71

Instruments: 1-3, Log(school_40-64); 4-6, Log(tertiary); 7-9, Log(school_40-64) and Log(tertiary). All regressions include time dummies. Standard errors are robust to heteroskedasticity and clustered at the country level. K-P is the Kleibergen-Paap weak identification test statistic. F-statistic refers to the Wald test on instruments in first-stage regressions. *t* statistics in parentheses. * p < 0.10, ** p < 0.05, *** p < 0.01.

		2010			1 11/1			CMM	
	(1)	25L5 (2)	(3)	(4)	LIML (5)	(6)	(7)	GMM (8)	(9)
Log(GDP)	-0.690***	-0.683***	-0.657***	-0.780***	-0.760***	-0.745***	-0.579***	-0.593***	-0.630***
	(-2.935)	(-2.837)	(-3.281)	(-2.774)	(-2.688)	(-3.073)	(-3.456)	(-3.209)	(-4.287)
Log(infant mortality)	-0.396*	-0.357	-0.327*	-0.467*	-0.419*	-0.392*	-0.310**	-0.297**	-0.361***
	(-1.741)	(-1.635)	(-1.680)	(-1.759)	(-1.668)	(-1.721)	(-2.199)	(-2.247)	(-3.132)
Log(schoolw 15-39)	0.518***	0.465***	0.475***	0.503***	0.452***	0.462***	0.533***	0.473***	0.378***
	(3.706)	(3.063)	(3.139)	(3.267)	(2.744)	(2.748)	(6.509)	(4.850)	(3.969)
Log(urbanization)		0.233*	0.176		0.232	0.173		0.215**	0.231***
		(1.730)	(1.372)		(1.612)	(1.260)		(2.025)	(2.601)
Log(abortion)			0.047			0.053			0.039
			(1.352)			(1.463)			(1.371)
Hansen J statistic	5.536	5.520	7.876	5.190	5.201	7.598	5.536	5.520	7.876
J p-value	0.853	0.854	0.641	0.878	0.877	0.668	0.853	0.854	0.641
K-P underidentification rk	19.142	19.464	17.262						
K-P p-value	0.059	0.053	0.100						
Anderson-Rubin F-statistic	67.221	5.276	5.607						
A-R p-value	0.000	0.000	0.000						
Observations	296	296	270	296	296	270	296	296	270
Countries	37	37	35	37	37	35	37	37	35

Table 9. Determinants of fertility transition onsets. Panel IV within estimates. Dependent variable: *onset*. Endogenous regressors: Log(GDP), Log(infant mortality), Log(school_15-39). Estimation methods: 2SLS, LIML, GMM

Instruments: predicted mortality and predicted mortality interacted with 2 dummies for the periods 1960-1964, 1965-1969, and with two dummies for Latin America and Middle East-North Africa. All regressions include time dummies. Standard errors are robust to heteroskedasticity and clustered at the country level. *t* statistics in parentheses. * *p* < 0.10, ** *p* < 0.05, *** *p* < 0.01

Table 10. Determinants of fertility transition onsets. First stage in 2SLS panel within estimates. Dependent variables: Log(*GDP*), Log(infant mortality), Log(school_15-39).

		(1)			(2)			(3)	
Equation	GDP	Infant mortality	Schooling	GDP	Infant mortality	Schooling	GDP	Infant mortality	Schooling
Log(Urbanization)				0.253 (0.790)	-0.245 (-1.148)	0.284 (1.228)	0.150 (0.548)	-0.205 (-1.051)	0.325 (1.528)
Log(Abortion)							0.087 [*] (1.870)	-0.022 (-0.611)	-0.027 (-0.578)
Log(School_40-64)	0.247 [*] (1.824)	-0.372*** (-4.221)	0.560*** (3.828)	0.190 (1.630)	-0.316*** (-4.068)	0.496*** (3.550)	0.194 (1.490)	-0.332*** (-4.828)	0.477 ^{***} (3.292)
Log(School_40-64)*D1960-1964	-0.049 (-1.541)	0.129*** (4.220)	-0.029 (-1.290)	-0.040 (-1.392)	0.121 ^{***} (4.909)	-0.019 (-1.000)	-0.029 (-0.846)	0.119 ^{***} (5.599)	-0.026 (-1.244)
Log(School_40-64)*D1970-1974	-0.041 (-0.788)	0.003 (0.086)	0.139 ^{***} (3.282)	-0.036 (-0.760)	-0.001 (-0.025)	0.144 ^{***} (3.450)	-0.019 (-0.407)	-0.013 (-0.356)	0.142*** (3.364)
World income	0.328 ^{***} (9.152)	-0.273**** (-7.693)	-0.047 (-1.180)	0.351*** (7.536)	-0.296*** (-7.239)	-0.022 (-0.523)	0.316*** (9.064)	-0.271*** (-9.328)	-0.012 (-0.311)
(World income)*South Asia&East Asia&Pacific	0.352*	-0.149	0.087	0.304	-0.102	0.033	0.380*	-0.126	0.007
(World income)*D1970-1974	(1.929) -0.002 (1.172)	(-1.491) 0.002** (2.163)	-0.000 (0.100)	-0.002	(-0.866) 0.002** (2.112)	-0.000 (0.278)	(1.897) -0.003* (1.803)	(-1.211) 0.003^{**} (2.242)	0.000
Predicted mortality	(-1.172) -2.888*** (-3.474)	(2.103) -2.376*** (-3.509)	(-0.190) 0.959 (0.739)	(-1.109) -2.797*** (-3.114)	(2.112) -2.464*** (-2.952)	(-0.278) 1.062 (1.082)	(-1.803) -3.053*** (-3.444)	(2.342) -2.508*** (-2.974)	(0.141) 1.147 (1.184)
Predicted mortality*Latin America	0.122 (1.091)	0.237 ^{**} (2.371)	0.093 (0.748)	0.127 (1.025)	0.232 ^{**} (2.290)	0.099 (0.867)	0.181 (1.467)	0.232*** (2.835)	0.039 (0.319)
Predicted mortality *Middle East& North Africa	-0.132	0.621***	-1.512***	-0.144	0.632***	-1.525***	-0.115	0.615***	-1.554***
	(-0.466)	(2.771)	(-3.605)	(-0.543)	(3.137)	(-3.694)	(-0.488)	(3.201)	(-3.839)
Predicted mortality* Sub-Saharan Africa	-0.113	0.332***	0.182^{*}	-0.119	0.338***	0.175^{*}	-0.175	0.356***	0.200^{**}
	(-1.039)	(5.012)	(1.686)	(-1.106)	(5.361)	(1.919)	(-1.481)	(5.830)	(2.040)

Predicted mortality *Europe&Central Asia	-0.361***	0.451***	-0.474***	-0.302	0.394***	-0.408***	-0.399**	0.418***	-0.375**
	(-2.871)	(5.140)	(-3.564)	(-1.533)	(3.006)	(-2.625)	(-2.275)	(3.582)	(-2.438)
Predicted mortality*D1960-1964	2.963***	2.156***	-1.362	2.898****	2.219***	-1.436	3.109***	2.282***	-1.515*
-	(3.837)	(3.293)	(-1.135)	(3.514)	(2.786)	(-1.578)	(3.807)	(2.823)	(-1.695)
Predicted mortality*D1965-1969	2.905***	2.253***	-1.213	2.834***	2.322***	-1.293	3.045***	2.377***	-1.364
	(3.717)	(3.444)	(-1.003)	(3.405)	(2.901)	(-1.416)	(3.668)	(2.926)	(-1.523)
Sanderson-Windmeijer F-statistic	10.690	10.502	172.772	6.175	8.567	73.996	6.776	10.545	66.924

Instruments: predicted mortality and predicted mortality interacted with 2 dummies for the periods 1960-1964, 1965-1969, and with two dummies for Latin America and Middle East-North Africa. All regressions include time dummies. Standard errors are robust to heteroskedasticity and clustered at the country level. *t* statistics in parentheses. * *p* < 0.10, ** *p* < 0.05, *** *p* < 0.01.

Table 11. Cross-country correlation matrix

	Observations	Transition	GDP	Infant mortality	Schoolw_15- 39	Urbanization	Catholic	Muslim	Women vote	Abortion
Transition	75	1								
GDP	75	0.206	1							
Infant mortality	75	-0.723	-0.526	1						
Schoolw_15-39	75	0.661	0.191	-0.708	1					
Urbanization	75	0.377	0.700	-0.567	0.193	1				
Catholic	73	0.160	0.048	-0.218	0.369	0.221	1			
Muslim	73	-0.113	0.208	0.082	-0.469	0.232	-0.528	1		
Women vote	73	-0.344	0.377	0.075	-0.322	0.122	-0.253	0.378	1	
Abortion	71	0.139	-0.007	-0.078	0.157	0.091	-0.202	-0.084	-0.094	1

Table 12. Determinants of fertility transition onsets. Dependent variable: transition.Lewbel's (2012) estimates

	(1)	(2)	(3)	(4)
GDP	-0.011* (-1.783) [-0.032]	-0.012* (-1.939) [-0.035]	0.001 (0.190) [0.004]	-0.004 (-0.664) [-0.012]
<i>schoolw</i> 15 – 39	0.042** (2.633) [0.109]	0.048 ^{***} (2.838) [0.122]	0.044 ^{**} (2.520) [0.114]	0.059*** (3.676) [0.150]
log(infant mortality)	-0.222*** (-3.805)	-0.210*** (-3.402)	-0.132*** (-2.667)	-0.122** (-2.376)
log(urbanization)	0.051* (1.786)	0.064** (2.191)	0.045 (1.578)	0.071*** (2.844)
South Asia	0.065* (1.748)	0.055 (1.500)	0.075 (1.558)	0.071 (1.490)
Middle East & North Africa	-0.001 (-0.028)	-0.016 (-0.352)	0.036 (0.780)	0.023 (0.523)
Muslim*East Asia&Pacific	0.203 (1.315)	0.172 (1.125)	0.239** (2.652)	0.215 [*] (1.719)
Catholic		-0.040 (-0.691)		
Women vote			-0.003*** (-3.146)	-0.002*** (-2.704)
log(abortion)				0.005 (0.348)
Hansen J	5.705	6.176	5.673	7.372
J p-value	0.222	0.186	0.225	0.117
K-P underidentification rk LM statistic	21.665	22.127	24.449	20.880
K-P p-value	0.001	0.000	0.000	0.001
Breusch-Pagan test. Equation GDP	38.793	58.857	68.200	64.478
B-P p-value	0.000	0.000	0.000	0.000
Breusch-Pagan test. Equation Schooling	8.388	10.143	10.264	10.330
B-P p-value	0.078	0.071	0.068	0.111

Breusch-Pagan test. Equation Mortality	7.479	11.501	9.616	13.285
B-P p-value	0.113	0.042	0.087	0.039
Sanderson-Windmeijer F-stat. Eq. GDP	102.220	88.744	64.870	45.051
Sanderson-Windmeijer F-stat. Eq. Schooling	92.687	80.232	108.447	83.734
Sanderson-Windmeijer F-stat. Eq. Mortality	45.645	44.819	56.214	23.609
Anderson-Rubin F-statistic	13.912	15.147	9.011	8.888
A-R statistic p-value	0.000	0.000	0.000	0.000
RESET	0.836	0.924	0.379	0.052
RESET p-value	0.361	0.337	0.538	0.820
Observations	72	72	72	70

t statistics in parentheses * p < 0.10, ** p < 0.05, *** p < 0.01