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Summary

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Keywords: Fertility, Demographic Trends, Female Education

JEL Classification: J11, J13, I15, I25, C26

This paper was presented at the 32nd Annual ESPE Conference in Antwerp on June 2018, and at the 2nd International Conference on Globalization and Development, 2018, in Göttingen.

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

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24th April 2019

Abstract

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

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1 Introduction

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

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

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

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

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

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

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

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

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

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

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

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

2 Related Literature

Several scholars of the fertility transition, demographers in particular, have stressed the importance of improved health status and declining mortality. The fact that, historically in several countries, mortality reductions preceded that of births provides a strong argument in support of this hypothesis (Kirk, 1996). The relationship between infant and child mortality and the choice of the number of children is the focus of a significant strand of the theoretical literature (e.g., Ben-Porath, 1976; Sah, 1991; Ehrlich and Lui, 1991; Boldrin and Jones, 2002; Kalemli-Ozcan, 2002; Soares, 2005; Doepke, 2005). In high-mortality environments, risk-averse parents may try to insure themselves against child death with more births. In this setting, a decline in infant mortality decreases uncertainty over child survival and reduces the demand for children. If the family's decision problem follows the sequence of births and deaths in the reproductive age, parents have an optimal desired number of surviving children and replace those dying in the early years of life with other births, until the targeted family size is reached. According to the replacement effect, decreasing infant mortality reduces total fertility without changing net fertility.

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

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

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

[3] Galor (2011) assesses this approach.

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

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

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

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

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

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

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

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

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

3 Modelling Fertility-Rate Transition Paths and Convergence

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

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

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

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

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

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

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

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

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

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

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

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

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 left-hand side of the regression equation (5) diverges to $-\infty$. We perform a one-sided test of the null hypothesis of convergence ($\gamma\geqslant 0$)) using the t-statistic calculated using the estimate of γ and a heteroskedasticity- and autocorrelation-consistent (HAC) standard error. This statistic is asymptotically distributed as a standard normal. Appendix A provides more detail on the *log t* test.

Phillips and Sul (2007) propose a procedure to identify clusters of units where convergence occurs in the long run. The algorithm starts with the ordering of all units in the panel according to the value of the last available observation, X_{iT}. Then, a core subgroup of units can be detected and the *log t* test can be used to assess whether another unit belongs to the core group or not. This procedure can be iterated until every unit in the panel is classified. In the following subsection, we apply this clustering method to the fertility rates in a panel of 180 countries.

3.1 Group Fertility-Rate Convergence

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

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

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

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

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

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

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

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

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

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

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

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

But this is not the case for the other countries we consider in our convergence analysis. Indeed, the fertility paths of many countries of Group I and Group H-L are characterized by an initially high and increasing TFR followed by a decreasing path that leads to different TFR values. The countries of Group H-L exhibit a stronger change in reproductive behavior. Indeed, Figure 1 shows that in this group, on average, the onset of the fertility decline occurs during the sixties, before other transitioning countries, and then it proceeds along a steep path. At the end of the period, average Group H-L TFR approaches that of Group L-L. The curve depicting the dynamics of TFR of countries in Groups I displays the same shape as that of Group H-L, but the declining part of the curve has a more gradual slope and at the end of the period fertility is still significantly higher than that of Group L.

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

4 The Onset Year of Fertility Transitions

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

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

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

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

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

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

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

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

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

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

5 Determinants of Fertility Transition Onsets

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

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

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

5.1 Econometric Model and Data

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

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

where GDP is the real per-capita GDP, $infant\ mortality$ is the infant mortality rate, $schoolw\_15-39$ denotes the average years of primary school of women in the 15–39 age range. Several studies find that primary education is particularly effective in explaining fertility (e.g., Murtin, 2013). The variable $schoolw\_15-39$ stresses the role of women's human capital in the choice of fertility (Galor and Weil, 1996). $\mathbf{x}$ is a vector of time-varying and time-invariant control variables, μ_t is the time effect, c_i is the country effect, which can be assumed correlated or uncorrelated with the random error u_{it}. The time index t refers to nine five-year nonoverlapping periods from 1955–1959 to 1995–1999.[10] In particular, the variable *onset* has been constructed as described above but with reference to five-year periods. Indeed, the benchmark period is defined by the interval that includes the year of the onset estimated in Section 4. Hence, *onset* takes value zero if t precedes the time interval during which a transition started, and value one otherwise. In this way, measurement error in the dependent variable *onset* should be smaller. The time effects imply that our model accounts for nonlinear trends common to all countries in the sample.

5.1.1 Data

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

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

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

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

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

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

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

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

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

5.1.2 Choice of Instrumental Variables

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

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

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

5.2 Results

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

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

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

5.2.1 IV Estimation: One Endogenous Regressor

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

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

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

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

The last application of the one-endogenous-regressor approach refers to women's educational attainment. Table 7 presents the results of six regressions: In the first three, we instrument log *schoolw_ 15–39* with log *school_ 40–64*, while in the last three, we use two IV: log *school_ 40–64* and log *school_ 40–64*, interacted with a dummy variable for the South Asian countries. In every specification the coefficients of log *schoolw_ 15–39* and

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

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

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

5.2.2 IV Estimation: Multiple Endogenous Regressors

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

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

Hence, the set of instruments includes the variables: *world income*, plus two variables obtained from the interactions between *world income* and the dummies for the period 1970–1974 and for the countries of South Asia, East Asia and Pacific; log *school_40–64*, and its product with the dummies for the periods 1960-1964 and 1970-1974; *predicted mortality*, plus six variables obtained from the interactions between *predicted mortality* and the dummies for the periods 1960–1964 and 1965–1969, for Latin American countries, for Sub-Saharan African countries, for Europe and Central Asia, and for the Middle East and North Africa. The last six interactive instruments capture the distinctive impact of the international epidemiological transition on infant mortality in the 1960s and 1970s with respect to more recent decades, and account for regional differences in the same effect.

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

Table C2 in online Appendix C presents the estimates of the coefficients of the IV in the first-stage equations. In each equation, the IV specific of the endogenous regressor display significant coefficients with the expected signs. To assess the validity of the IV used in the second-stage estimates, we test for underidentification of the regression model, weak instruments, and overidentifying restrictions. In Table 8, all the regressions display values of the Kleibergen–Paap rk Lagrange-multiplier test statistic that suggest the rejection of the null of underidentification at the 10% significance level. We test for weak instruments using the conditional first-stage F-statistic Sanderson and Windmeijer (2016) recently proposed for equations with multiple endogenous variables. Assuming the error term is i.i.d, this statistic can be used to test the null of weak IV in relation to a given maximum relative bias. Sanderson and Windmeijer (2016) argue that the critical values tabulated in Stock and Yogo (2005) can be used to conduct the test with the conditional F-statistic. However, the assumption of i.i.d disturbances often seems unjustified, particularly in linear-probability models.[15] Accordingly, all the test statistics in our regressions are robust to heteroskedasticity. In this case, testing weak IV is usually conducted by applying Staiger and Stock's (1997) rule of thumb.

In our regressions, this rule suggest the reduced-form log *schoolw_ 15–39* is unaffected by the use of weak IV. Indeed, in this case the conditional F-statistic assumes values greater than 66. In the estimates of the reduced form of log *infant mortality*, the same statistic takes values greater than 10 in two specifications and close to 10 in the third (Sanderson-Windmeijer F-statistic=8.56). Identification of the reduced form log *GDP* seems more problematic. Indeed, the Sanderson-Windmeijer F-statistic takes value 10.69 in one model specification, and 6.17, 6.77 in other two. However, it can be noted that in the second-stage equations the three estimates of the coefficient of log *GDP* are almost identical. Further support to our estimates comes from Anderson and Rubin's F test because in all regressions, we reject the null of the Anderson–Rubin test at the 1% level. The Hansen J statistic suggests the validity of the orthogonality conditions in all regressions.

Recent research on the IV estimation method (Young, 2018; Andrews, 2018) highlights the negative influence of outliers and non-homoskedastic errors on the quality of inference, in particular when weak IV are used. In online Appendix C2, we show that outliers have little influence on our GMM estimates of the basic model (specification 7 in Table 8), and the null of normally distributed residuals cannot be rejected. We estimate the model removing one country in each regression. Table C3, panel *A* shows how the estimates of the coefficients of the three endogenous regressors remain largely significant in all but one sample that gives a p-value of 0.105 for the coefficient of log *infant mortality*. Similar, but less strong, evidence comes from panel *B* in Table C3 that summarizes the distribution of the estimate of the first-stage F-statistic. The equation of log *schoolw_ 15–39* remains always strongly identified, while 66% of the Fs from the estimates of log *GDP* and log

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

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

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

5.2.3 Identification through Heteroskedasticity

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

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

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

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

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

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

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

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

the usual significance levels.

To compare the marginal effect of the logged variables with that of the untransformed variables, Table 9 presents in square brackets the semielasticity, the change of *transition* after a proportionate increase in the regressor, averaged across the countries in the sample. In the first two specifications the coefficient of *GDP* is statistically significant but in the other two the null that the coefficient is zero cannot be rejected. In every specification the coefficients of *schoolw_ 15–39* and log *infant mortality* are strongly significant. In the most complete specification, the coefficient of *schoolw_ 15–39* implies that if the educational attainment of women increases by 10% the year of the onset comes 10.17 months earlier, while the corresponding coefficients for infant mortality implies the event occurs 7.19 months earlier.[18]

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

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

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

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

6 Conclusions

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

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

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

References

Acemoglu, Daron, and Simon Johnson, "Disease and Development: the Effect of Life Expectancy on Economic Growth," *Journal of Political Economy* 115:6 (2007), 925-985.

Acemoglu, Daron, Simon Johnson, James A. Robinson, and Pierre Yared, "Income and Democracy," *The American Economic Review* 98:3 (2008), 808–42.

Andrews, Isaiah, James Stock, and Liyang Sun, "Weak Instruments in IV Regression:

Theory and Practice," Harvard University, Mimeo (2018).

Angrist, Joshua D., and Jörn-Steffen Pischke, "*Mostly Harmless Econometrics: An Empiricist's Companion*," (Princeton: Princeton University Press, 2009).

Barro, Robert J., "Convergence and Modernisation," *The Economic Journal* 125:585 (2015), 911-942.

Barro, Robert J., and Jong Wha Lee, "A New Data Set of Educational Attainment in the World, 1950–2010," *Journal of Development Economics* 104 (2013), 184-198.

Barro, Robert J., and Jong Wha Lee, "*Education Matters: Global Schooling Gains from the 19th to the 21st Century*," (New York: Oxford University Press, 2015).

Barro, Robert J., and Rachel M. McCleary, "Religion and Economic Growth across Countries," *American Sociological Review* 68:5 (2003), 760-781.

Baum, Christopher F., Mark E. Schaffer, and Steven Stillman, "Enhanced Routines for Instrumental Variables/GMM Estimation and Testing," *Stata Journal* 7:4 (2007), 465-506.

Baum, Christopher F., Mark E. Schaffer, and Steven Stillman, "ivreg2: Stata Module for Extended Instrumental Variables/2SLS, GMM and AC/HAC, LIML and K-class Regression," (2010). http://ideas.repec.org/c/boc/bocode/s425401.html

Bazzi, Samuel, and Michael A. Clemens, "Blunt Instruments: Avoiding Common Pitfalls in Identifying the Causes of Economic Growth," *American Economic Journal: Macroeconomics* 5:2 (2013), 152-86.

Becker, Gary S, "An Economic Analysis of Fertility," in *Demographic and economic change in developed countries: a conference of the Universities-National Bureau Committee for Economic Research* (New York: Columbia University Press).

Becker, Gary S., and H. Gregg Lewis, "On the Interaction between the Quantity and Quality of Children," *Journal of Political Economy* 81:2, Part 2 (1973), S279-S288.

Ben-Porath, Yoram, "Fertility Response to Child Mortality: Micro Data from Israel," *Journal of Political Economy* 84:4, Part 2 (1976), S163-S178.

Black, Dan, A., Natalia Kolesnikova, Seth, G. Sanders, and Lowell, J. Taylor, "Are Children 'Normal'?," *Review of Economics and Statistics* 95:1 (2013), 21–33.

Bloom, David E., David Canning, Gunther Fink, and Jocelyn E. Finlay, "Fertility, Female Labor Force Participation, and the Demographic Dividend," *Journal of Economic Growth* 14:2 (2009), 79–101.

Blundell, Richard, and Stephen Bond, "Initial Conditions and Moment Restrictions in Dynamic Panel Data Models," *Journal of Econometrics* 87:1 (1998), 115-143.

Boldrin, Michele, and Larry E. Jones, "Mortality, Fertility, and Saving in a Malthusian Economy," *Review of Economic Dynamics* 5:4 (2002), 775-814.

Bongaarts, John, and Susan Cotts Watkins, "Social Interactions and Contemporary Fertility Transitions," *Population and Development Review* 22:4(1996), 639-682.

Brueckner, Markus, and Hannes Schwandt, "Income and Population Growth," *The Economic Journal* 125:589 (2015), 1653-1676.

Bun, Maurice J.G., and Frank Windmeijer, "The Weak Instrument Problem of the System GMM Estimator in Dynamic Panel Data Models," *The Econometrics Journal* 13:1 (2010), 95-126.

Cervellati, Matteo, and Uwe Sunde, "Life Expectancy and Economic Growth: the Role of the Demographic Transition," *Journal of Economic Growth* 16:2 (2011), 99-133.

Cervellati, Matteo, and Uwe Sunde, "The Economic and Demographic Transition, Mortality, and Comparative Development," *American Economic Journal: Macroeconomics* 7:3 (2015), 189-225.

Chesnais, Jean-Claude, *The Demographic Transition: Stages, Patterns, and Economic Implications* (Oxford: Oxford University Press, 1992).

Delventhal, Matthew J., Jesús Fernández-Villaverde, and Nezih Guner, "Demographic Transitions Across Time and Space," Brown University,Working Paper, December, 2018.

Diebolt, Claude, and Faustine Perrin, "*From Stagnation to Sustained Growth: the Role of Female Empowerment,*" American Economic Review 103:3 (2013) 545-49.

Doepke, Matthias, "Child Mortality and Fertility Decline: Does the Barro-Becker Model Fit the Facts?," *Journal of Population Economics* 18:2 (2005), 337-366.

Dorius, Shawn F., "Global Demographic Convergence? A Reconsideration of Changing Intercountry Inequality in Fertility," *Population and Development Review* 34:3 (2008), 519-537.

Emran, M. Shahe, and Zhaoyang Hou, "Access to Markets and Rural Poverty: Evidence from Household Consumption in China," *Review of Economics and Statistics* 95:2 (2013), 682-697.

Ehrlich, Isaac, and Francis T. Lui, "Intergenerational Trade, Longevity, and Economic Growth," *Journal of Political Economy* 99:5 (1991), 1029-1059.

Galor, Oded, *Unified growth theory* (Princeton: Princeton University Press, 2011).

Galor, Oded, and David N. Weil, "The Gender Gap, Fertility, and Growth," *The American Economic Review* 86:3 (1996), 374-387.

Galor, Oded, and David N. Weil, "Population, Technology, and Growth: From Malthusian Stagnation to the Demographic Transition and Beyond," *American Economic Review* 90:4 (2000), 806-828.

Greenwood, Jeremy, Nezih Guner, and Guillaume Vandenbroucke, "Family Economics Writ Large," *Journal of Economic Literature* 55:4 (2017), 1346–143

Hansen, Casper Worm, and Lars Lønstrup, "The Rise in Life Expectancy and Economic Growth in the 20th Century," *Economic Journal* 125:584 (2015), 838-852.

Hausmann, Ricardo, Lant Pritchett, and Dani Rodrik, "Growth Accelerations," *Journal of Economic Growth* 10:4 (2005), 303-329.

Herzer, Dierk, Holger Strulik, and Sebastian Vollmer, "The Long-Run Determinants of Fertility: one Century of Demographic Change 1900–1999," *Journal of Economic Growth* 17:4 (2012), 357-385.

Jones, Benjamin F., and Benjamin A. Olken, "The Anatomy of Start-Stop Growth," *The Review of Economics and Statistics* 90:3 (2008), 582-587.

Kalemli-Ozcan, Sebnem, "Does the Mortality Decline Promote Economic Growth?," *Journal of Economic Growth* 7:4 (2002), 411-439.

Kirk, Dudley, "Demographic Transition Theory," *Population Studies* 50:3 (1996), 361-387.

Kiszewski, Anthony, Andrew Mellinger, Andrew Spielman, Pia Malaney, Pia, Sonia Ehrlich Sachs, and Jefrey Sachs, "A Global Index Representing the Stability of Malaria

Transmission," *The American Journal of Tropical Medicine and Hygiene* 70:5 (2004), 486–498.

Kleibergen, Frank, and Richard Paap, "Generalized Reduced Rank Tests Using the Singular Value Decomposition," *Journal of Econometrics* 133:1 (2006), 97-126.

Lewbel, Arthur, "Using Heteroscedasticity to Identify and Estimate Mismeasured and Endogenous Regressor Models," *Journal of Business & Economic Statistics* 30:1 (2012), 67-80.

Lindo, Jason M, "Are Children Really Inferior Goods? Evidence from Displacement-Driven Income Shocks," *Journal of Human Resources* 45:2 (2010), 301-327.

Lorentzen, Peter, John McMillan, and Romain Wacziarg, "Death and Development," *Journal of Economic Growth* 13:2 (2008), 81-124.

McCord, Gordon C., Dalton Conley, and Jeffrey D. Sachs, "Malaria Ecology, Child Mortality and Fertility," *Economics and Human Biology* 24:1 (2017), 1-17.

Millimet, Daniel L., and Jayjit Roy, "Empirical Tests of the Pollution Haven Hypothesis when Environmental Regulation is Endogenous," *Journal of Applied Econometrics* 31:4 (2016), 652-677.

Murtin, Fabrice, "Long-Term Determinants of the Demographic Transition, 1870–2000," *Review of Economics and Statistics* 95:2 (2013), 617-631.

Olea, José Luis Montiel, and Carolin Pflueger, "A Robust Test for Weak Instruments," *Journal of Business & Economic Statistics* 31:3 (2013), 358-369.

Perron, Pierre, and Xiaokang Zhu, "Structural Breaks with Deterministic and Stochastic Trends," *Journal of Econometrics* 129:1-2 (2005), 65-119.

Perron, Pierre, and Tomoyoshi Yabu, "Testing for Shifts in Trend with an Integrated or Stationary Noise Component," *Journal of Business & Economic Statistics* 27:3 (2009), 369-396.

Pesaran Hashem M., and Larry W. Taylor, "Diagnostics for IV regressions," *Oxford Bulletin of Economics and Statistics* 61:2 (1999), 255-281.

Phillips, Peter C. B., and Donggyu Sul, "Transition Modeling and Econometric Convergence Tests," *Econometrica* 75:6 (2007), 1771-1855.

Phillips, Peter C. B., and Donggyu Sul, "Economic Transition and Growth," *Journal of Applied Econometrics* 24:7 (2009), 1153-1185.

Reher, David S., "The demographic transition revisited as a global process," *Population, Space and Place* 10.1 (2004), 19-41.

Sah, Raaj K., "The Effects of Child Mortality Changes on Fertility Choice and Parental Welfare," *Journal of Political Economy* 99:3 (1991), 582-606.

Sanderson, Eleanor, and Frank Windmeijer, "A Weak Instrument F-Test in Linear IV Models with Multiple Endogenous Variables," *Journal of Econometrics* 190:2 (2016), 212-221.

Schultz, T. Paul, "Demand for Children in Low Income Countries," in Rosenzweig, Mark R., and Oded Stark eds., *Handbook of Population and Family Economics vol. 1a.* (Amsterdam: Elsevier Science, 1997).

Scotese Lehr, Carol, "Evidence on the Demographic Transition," *The Review of Economics and Statistics* 91:4 (2009), 871-887.

Soares, Rodrigo R., "Mortality Reductions, Educational Attainment, and Fertility Choice," *American Economic Review* 95:3 (2005), 580-601.

Staiger, Douglas, and James H. Stock, "Instrumental Variables Regression with Weak Instruments," *Econometrica* 65:3 (1997), 557-586.

Stock, James H., and Motohiro Yogo, "Testing for Weak Instruments in Linear IV Regression," in Donald W. K. Andrews and James H. Stock eds., *Identification and Inference for Econometric Models: Essays in Honor of Thomas Rothenberg* (Cambridge: Cambridge University Press, 2005).

Strulik, Holger. and Sebastian Vollmer, "The Fertility Transition Around the World 1950-2005," *Journal of Population Economics* 28:1 (2015), 31-44.

Young, Alwyn, "Consistency Without Inference: Instrumental Variables in Practical Application," Unpublished manuscript, London: LSE, (2018).

Table 1. Descriptive statistics of fertility (average TFR) in convergence groups

	Countries	Min 1950	Min 2015	Max 1950	Max 2015	Mean 1950	Mean 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. Descriptive statistics. Sample of 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 per capita	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

Table 3. Correlation matrix

Correlation Matrix	Obs.	Onset	Log GDP p. c.	Log Infant mortal.	Log schoolw_15-39	Log urbanization	Log abortion
Onset	675	1					
Log(GDP per capita)	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 4. Determinants of fertility transition onsets. Dependent variable: *onset*.
Panel fixed and random effects estimates

	(1)	(2)	(3)	(4)	(5)	(6)
	Fixed Effects		Random Effects			
Log(GDP per capita)	-0.202***	-0.163**	-0.196***	-0.169***	-0.163***	-0.138***
	(0.050)	(0.070)	(0.037)	(0.042)	(0.042)	(0.048)
Log(infant mortality)	-0.141***	-0.004	-0.153**	-0.117*	-0.117*	0.008
	(0.053)	(0.069)	(0.067)	(0.068)	(0.070)	(0.076)
Log(schoolw_15-39)	0.244***	0.335***	0.230***	0.258***	0.256***	0.335***
	(0.049)	(0.047)	(0.043)	(0.049)	(0.050)	(0.051)
Log(urbanization)	0.435***	0.436***	0.264***	0.247***	0.236***	0.240***
	(0.043)	(0.051)	(0.043)	(0.052)	(0.051)	(0.060)
catholic				-0.001	-0.001	-0.001
				(0.001)	(0.001)	(0.001)
muslim				0.001	0.001	0.001
				(0.001)	(0.001)	(0.001)
women vote					-0.001	0.000
					(0.001)	(0.001)
Log(abortion)		0.023				0.022
		(0.025)				(0.020)
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

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

Table 5. Determinants of fertility transition onsets. Panel IV within estimates. Dependent variable: *onset*. Endogenous regressor: GDP per capita

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Log(GDP per capita)	0.085	-0.053	0.115	0.527*	0.138	0.090	-0.545***	-0.383**	-0.357**
	(0.165)	(0.174)	(0.112)	(0.320)	(0.175)	(0.172)	(0.142)	(0.171)	(0.157)
Log(urbanization)		0.613***	0.591***		0.656***	0.688***		0.547***	0.532***
		(0.112)	(0.124)		(0.127)	(0.138)		(0.130)	(0.134)
Log(abortion)			-0.050			0.011			-0.010
			(0.037)			(0.047)			(0.030)
First stage									
oil shock	2.642***	2.661***	2.369***				0.950	0.941	1.318**
	(0.635)	(0.643)	(0.885)				(0.625)	(0.621)	(0.529)
world income				0.540***	0.446***	0.415***	0.403***	0.406***	0.407***
				(0.097)	(0.055)	(0.056)	(0.064)	(0.073)	(0.077)
Hansen J							0.577	0.195	0.308
J p-value							0.447	0.659	0.579
K-P Wald test F	17.300	17.124	7.165	31.054	66.645	55.580	20.805	17.001	19.176
Anderson-Rubin F-statistic	0.253	0.092	1.059	1.765	0.601	0.274	4.548	1.736	1.666
A-R p-value	0.617	0.763	0.308	0.189	0.441	0.603	0.015	0.186	0.199
Observations	426	426	384	594	594	478	376	376	346
Countries	64	64	61	66	66	64	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. The Anderson and Rubin's F-statistic tests the null hypothesis that the endogenous regressor is irrelevant. Standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table 6. Determinants of fertility transition onsets. Panel IV within estimates. Dependent variable: *onset*. Endogenous regressor: infant mortality

	(1)	(2)	(3)	(4)	(5)	(6)
Log(infant mortality)	-0.588***	-0.561**	-0.642**	-0.587***	-0.538**	-0.611**
	(0.113)	(0.254)	(0.269)	(0.113)	(0.251)	(0.268)
Log(urbanization)		0.067	0.017		0.096	0.055
		(0.440)	(0.441)		(0.432)	(0.431)
Log(abortion)			-0.064			-0.061
			(0.065)			(0.065)
First stage						
predicted mortality	1.243***	0.806***	0.811***	1.242***	0.793***	0.794***
	(0.132)	(0.188)	(0.179)	(0.138)	(0.194)	(0.185)
*predicted mortality*dummy SSA*				0.027	0.188	0.202
				(0.191)	(0.173)	(0.150)
Hansen J				0.923	0.957	1.019
J p-value				0.337	0.328	0.313
K-P Wald F-statistic	88.659	18.430	20.520	64.249	18.214	25.426
Anderson-Rubin F-statistic	40.003	9.934	11.900	22.420	5.983	7.501
A-R p-value	0.000	0.003	0.001	0.000	0.005	0.002
Observations	336	336	304	336	336	304
Countries	42	42	40	42	42	40

Instruments: *1-3, predicted mortality; 4-6, predicted mortality and predicted mortality interacted with dummy for 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. The Anderson and Rubin's F-statistic tests the null hypothesis that the endogenous regressor is irrelevant. Standard errors 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: Average years of primary school, female aged 15-39

	(1)	(2)	(3)	(4)	(5)	(6)
Log(schoolw_15-39)	0.470***	0.350***	0.353***	0.460***	0.332***	0.361***
	(0.041)	(0.054)	(0.054)	(0.046)	(0.062)	(0.054)
Log(urbanization)		0.313***	0.355***		0.332***	0.347***
		(0.066)	(0.073)		(0.073)	(0.073)
Log(abortion)			0.006			0.005
			(0.027)			(0.027)
First stage						
Log(school_40-64)	0.871***	0.724***	0.710***	0.844***	0.698***	0.702***
	(0.090)	(0.098)	(0.111)	(0.095)	(0.099)	(0.111)
*Log(school_40-64)*dummy South Asia*				0.259**	0.251***	0.286***
				(0.105)	(0.080)	(0.088)
Hansen J				1.247	1.324	0.912
J 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
Anderson-Rubin F-statistic	87.276	35.158	26.043	46.138	20.548	14.229
A-R p-value	0.000	0.000	0.000	0.000	0.000	0.000
Observations	674	674	518	674	674	518
Countries	75	75	71	75	75	71

Instruments: 1-3, Log(school_40-64); 4-6, Log(school_40-64) and Log(school_40-64) interacted with a dummy for countries of South Asia. 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. The Anderson and Rubin's F-statistic tests the null hypothesis that the endogenous regressor is irrelevant. Standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table 8. Determinants of fertility transition onsets. Panel IV within estimates. Dependent variable: *onset*.
Endogenous regressors: Log(GDP per capita), Log(infant mortality), Log(schoolw_15-39)

	2SLS			LIML			GMM		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Log(GDP per capita)	-0.690***	-0.683***	-0.657***	-0.780***	-0.760***	-0.745***	-0.579***	-0.593***	-0.630***
	(0.235)	(0.241)	(0.200)	(0.281)	(0.283)	(0.242)	(0.168)	(0.185)	(0.147)
Log(infant mortality)	-0.396*	-0.357	-0.327*	-0.467*	-0.419*	-0.392*	-0.310**	-0.297**	-0.361***
	(0.227)	(0.218)	(0.194)	(0.265)	(0.251)	(0.228)	(0.141)	(0.132)	(0.115)
Log(schoolw_15-39)	0.518***	0.465***	0.475***	0.503***	0.452***	0.462***	0.533***	0.473***	0.378***
	(0.140)	(0.152)	(0.151)	(0.154)	(0.165)	(0.168)	(0.082)	(0.098)	(0.095)
Log(urbanization)		0.233*	0.176		0.232	0.173		0.215**	0.231***
		(0.135)	(0.128)		(0.144)	(0.137)		(0.106)	(0.089)
Log(abortion)			0.047			0.053			0.039
			(0.035)			(0.036)			(0.029)
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 LM	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						
S-W test F-stat. Eq. GDP	10.690	6.175	6.776						
S-W test F-stat. Eq. Infant Mortality	10.502	8.567	10.545						
S-W test F-stat. Eq. Schooling	172.772	73.996	66.924						
Observations	296	296	270	296	296	270	296	296	270
Countries	37	37	35	37	37	35	37	37	35

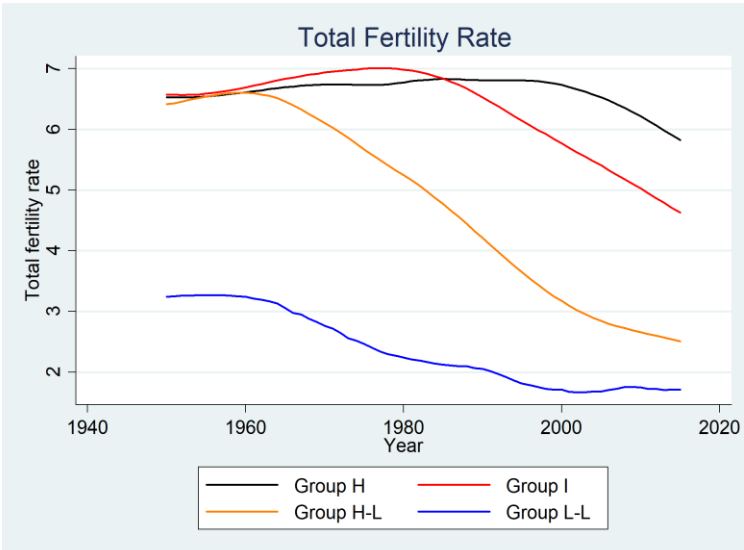
Instruments: world income plus two variables obtained from the interactions between world income and the dummies for the period 1970--1974 and for the countries of South Asia, East Asia and Pacific; predicted mortality plus six variables obtained from the interactions between predicted mortality and 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; log school_40--64, and its product with the dummies for the period 1960-1964 and 1970-1974.
All regressions include time dummies. Standard errors are robust to heteroskedasticity and clustered at the country level. K-P is the Kleibergen-Paap underidentification rk LM test-statistic. S-W is the Sanderson-Windmeijer test for weak instruments. The Anderson and Rubin's F-statistic tests the null hypothesis that the endogenous regressors are irrelevant. Standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table 9. Determinants of fertility transition onsets. Lewbel's (2012) estimates.
Dependent variable: *transition*.

	(1)	(2)	(3)	(4)
$\overline{GDP\ per\ capita}$	-0.011*	-0.012*	0.001	-0.004
	(0.006)	(0.006)	(0.007)	(0.006)
	[-0.032*]	[-0.035*]	[0.004]	[-0.010]
$\overline{schoolw\ 15-39}$	0.041**	0.047***	0.046**	0.062***
	(0.016)	(0.017)	(0.018)	(0.017)
	[0.106**]	[0.120***]	[0.117**]	[0.157***]
$\overline{\log(infant\ mortality)}$	-0.224***	-0.211***	-0.125***	-0.111**
	(0.060)	(0.063)	(0.049)	(0.052)
$\overline{\log(urbanization)}$	0.051*	0.064**	0.046	0.073***
	(0.029)	(0.030)	(0.029)	(0.025)
South Asia	0.064*	0.055	0.080	0.079
	(0.038)	(0.038)	(0.049)	(0.049)
Middle East & North Africa	-0.002	-0.017	0.037	0.025
	(0.045)	(0.046)	(0.046)	(0.044)
$\overline{Muslim}$*East Asia&Pacific	0.002	0.002	0.002***	0.002*
	(0.002)	(0.002)	(0.001)	(0.001)
$\overline{Catholic}$		-0.000		
		(0.001)		
Women vote			-0.003***	-0.002***
			(0.001)	(0.001)
$\overline{\log(abortion)}$				0.005
				(0.015)
Hansen J	5.885	6.338	5.5372	7.176
J p-value	0.208	0.175	0.237	0.127
K-P underidentification rk LM statistic	20.456	20.657	21.352	19.904
K-P p-value	0.001	0.001	0.001	0.001
Breusch-Pagan test. Equation GDP, $\chi^2(3)$	49.117	71.417	91.518	86.017
B-P p-value	0.000	0.000	0.000	0.000
Breusch-Pagan test. Equation Schooling, $\chi^2(3)$	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, $\chi^2(3)$	7.117	11.873	10.458	12.771
B-P p-value	0.127	0.037	0.063	0.047
Sanderson-Windmeijer F-stat. Eq. GDP	78.357	63.345	47.473	37.937
Sanderson-Windmeijer F-stat. Eq. Schooling	54.266	55.073	50.433	37.154
Sanderson-Windmeijer F-stat. Eq. Mortality	38.224	37.083	45.160	18.801
Anderson-Rubin F-statistic	11.702	12.326	8.707	8.699
A-R statistic p-value	0.000	0.000	0.000	0.000
RESET Wald test statistic, $\chi^2(1)$	0.708	0.784	0.695	0.365
RESET p-value	0.400	0.376	0.404	0.546
Observations	72	72	72	70

A bar over a variable denotes the average across the time dimension. Two-step efficient GMM estimation. Instruments: Lewbel's (2012) generated instruments, and the share in current GDP of merchandise export. Semielasticities in square brackets. K-P is the Kleibergen-Paap underidentification rk LM test-statistic. Breusch and Pagan is a test for disturbance homoscedasticity. The Sanderson-Windmeijer F-statistic tests for weak instruments. The Anderson-Rubin F-statistic tests the null hypothesis that the endogenous regressors are irrelevant. RESET is a Ramsey Wald $\chi^2(1)$ test for the null that there are no neglected nonlinearities. Standard errors in parentheses. * p < 0.10, ** p < 0.05, *** p < 0.01.

Figure 1. Fertility in Convergence Groups, 1950-2015



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

Appendix

December 9, 2019

A1 Phillips and Sul's (2007) approach to transition modeling

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:

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

where δ_i is fixed, ξ_{it} is $iid(0,1)$ across i and weakly dependent over t, and $L\left(t\right)$ is a slowly varying function[1], such that $L\left(t\right) \rightarrow \infty$ as $t \rightarrow \infty$. In model (A2), δ_{it} converges to δ_i for all $\alpha \geq 0$. Interestingly, the model

[1] A slowly varying function $F\left(t\right)$ satisfies the condition: $F\left(at\right)/F\left(t\right) \rightarrow 1$ as $t \rightarrow \infty$ for all $a > 0$.

accommodates heterogeneous transition paths both across individuals and over time[2].

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:

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

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:

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

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:

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

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:

$$H_0 : \delta_i = \delta \text{ and } \alpha \geq 0,$$

against the alternative:

$$H_A : \delta_i \neq \delta \text{ for all } i \text{ or } \alpha < 0.$$

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

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

[2] Phillips 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.

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.

A2 Convergence analysis of fertility rates

The application of the *log t* test to the time series of the full set of 180 countries leads to the rejection of the null hypothesis of convergence.[3] We use the clustering algorithm of Phillips and Sul (2007, 2009) to investigate the possibility of group convergence in TFR among countries. By ordering the 180 countries under consideration according to the last five years' average TFR, we find that the first country in the list is Niger with the highest fertility rate (7.37), while the last is Singapore (TFR=1.23). The whole procedure consists of a clustering algorithm and tests of group merging. In the first 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, $t_k > -1.65$).[4] In the second step of the algorithm 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. In the third step, we run the *log t* test for the group for which $\hat{t} < c$ and see if the null of convergence can be accepted (i.e., $t_b > -1.65$). If not, we repeat the first two steps on the remaining countries to determine whether this set can be subdivided in convergence groups. Phillips

[3] We estimate *log t* regressions with HAC standard errors using a Bartlett kernell and Newy-West fixed bandwidth selection procedure. Calculations were performed using the unofficial STATA command "hacreg" by Wang and Wu (2012).

[4] The core group size k^* is given by the criterion:

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

and Sul (2009) note that when sample size is small, the application of this conservative testing procedure could lead to more groups than the true number, and suggest the running of the *log t* test across subgroups. In the following, we summarize the results of the application of the clustering procedure to the TFR in the sample of this analysis.

First group. The core is formed by Niger, Somalia and the Democratic Republic of Congo ($t_k = 1.09$). The first cluster includes the core countries plus Chad and Timor-Leste: $t_b = 2.85$. The *log t* test on the remaining countries rejects the null: $t_b = -10.94$.

Second group. The core is formed by Mali and Burundi ($t_k = -0.30$), but no other country can be added to the core. The *log t* test on the remaining countries rejects the null: $t_b = -9.88$.

Third group. The core is formed by Angola and Uganda ($t_k = 4.79$). The third cluster includes the core countries plus Nigeria, Burkina Faso, Gambia, Equatorial Guinea: $t_b = 7.36$. The *log t* test on the remaining countries rejects the null: $t_b = -7.39$.

Fourth group. Repeating the clustering procedure again, we find the next country, Mozambique, does not satisfy the condition $\min\{t_k\} > -1.65$ ($t_k = -2.99$). The *log t* test on the remaining countries rejects the null: $t_b = -6.96$.

Fifth group. The core is formed by Tanzania, Afghanistan, Benin, Zambia, Cote d'Ivoire, Guinea, Central Africa Republic, Senegal, Cameroon, Malawi, Guinea Bissau ($t_k = 1.61$). The fifth cluster includes the core countries plus Mauritania, Congo, Sierra Leone, Liberia, Sudan, Togo, Sao Tomè and Principe, Comoros, Ethiopia, Iraq, Madagascar, Eritrea, Yemen, Rwanda: $t_b = -0.40$.

Sixth group. Running the *log t* test on the remaining countries, we accept the null of convergence: $t_b = -0.23$.

Next, we run the *log t* regression to test the hypothesis of group merging. We find that: the first group and the second can be merged: $t_b = 2.78$; Mozambique can be merged with the third group: $t_b = 3.75$; the first, second, and third groups and Mozambique can be merged: $t_b = -0.57$, we denote this cluster as Group H; Group H and the fifth group cannot be merged: $t_b = -20.94$; we denote the fifth group as Group I and the sixth group as Group L. Table A2 shows the composition of each convergence group.

B Perron and Yabu's (2009) test for structural change in the trend function

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

$$y_t = \mu + \beta_1 t + \beta_b B_t + u_t, \qquad t = 1, ..., T \tag{B1}$$

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

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

Note that the trend function in this model remains joined at the year of the break. Otherwise, the trend would verify a discontinuity at T_1. Perron and Yabu (2009) propose a test of the null hypothesis of one break in the slope of the trend function (B1). Under the null $\beta_b = 0$, meaning the trend function (B1) does not change in the period under consideration. The test procedure can be applied to models as (B1) 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 (B1) 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,

$$(1 - \alpha L)\, y_t = (1 - \alpha L)\, \mathbf{x}_t^{'} \Psi + (1 - \alpha L)\, u_t, \tag{B2}$$

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 (B2) requires an estimate of α that can be obtained from the OLS regression:

$$\widehat{u}_t = \alpha \widehat{u}_{t-1} + \sum_{i=1}^{k} \zeta_i \Delta \widehat{u}_{t-i} + e_t, \tag{B3}$$

where $\widehat{u}_t$ is the residual from the OLS estimation of model (B3). 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,

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

for $\delta \in (0,1)$ and $d > 0$. Perron and Yabu's (2009) Theorem 1 asserts that the Wald test obtained using $\widehat{\alpha}_S$ 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).

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 (B1) where the estimate of β_b is negative. As in Perron and Zhu (2005), the estimation of the coefficients of (B1) is limited to the range of the time series defined as $\varepsilon \leq \lambda_1 \leq 1 - \varepsilon$, where $\lambda_1 = T_1/T$, and ε is a trimming parameter. In this paper, we set ε to 10% and search for a break date between 1956 and 2009.

Table B1 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 B1 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 Figure B1. In each graph, a vertical line represents the estimate of the onset year. Table B1 does not show the results of Aruba and Singapore because the trend of TFR is always decreasing between 1956 and 2009, and the econometric procedure gives $T_1 = 2009$ in both cases.

C1 Variable description and data sources

Total Fertility Rate. Total fertility rate represents the number of children that would be born to a woman if she were to live to the end of her childbearing years and bear children in accordance with age-specific fertility rates of the specified year. The source of the annual data for the period 1960-2015 is the database World Development Indicators available online at: http://databank.worldbank.org/data/home.aspx. The source of the annual data for the period 1950-1959 is the database World Population Prospects: The 2017 Revision of the United Nations Population Division. These data are estimates produced on the basis of national statistical sources. Details can be found in the Methodology Report available at: https://population.un.org/wpp/Publications/Files/WPP2017_Methodology.pdf.

Real per-capita GDP. Real per-capita GDP in 1990 international Geary Khamis Dollars. Source: The Maddison Project 2013, Bolt and Van Zanden (2014), available at:

https://www.rug.nl/ggdc/historicaldevelopment/maddison/releases/maddison-project-database-2013.

Each observation is an average of annual data for each 5-year period.

Infant mortality rate. Probability of dying between birth and exact age 1. Source: United Nations, Department of Economic and Social Affairs, Population Division (2017). World Population Prospects: The 2017 Revision, DVD. These data are estimates produced on the basis of national statistical sources. Details can be found in the Methodology Report available at: https://population.un.org/wpp/Publications/Files/WPP2017_Methodology.pdf.

Schoolw_15-39. The average years of primary school of women in the 15-39 age range. Calculated from data on educational attainment and population disaggregated by sex and by 5-year age intervals. Source: Barro and Lee (2013), http://www.barrolee.com/.

Urbanization. Percentage of population at mid-year residing in urban areas. Source: United Nations, Department of Economic and Social Affairs, Population Division (2014). World Urbanization Prospects: The 2014 Revision, CD-ROM Edition.

Catholic and Muslim. The Catholic and Muslim fractions of population in 1970 and 2000. Data drawn from the Religion Adherence dataset used by Barro and McCleary (2003), and available at:

https://scholar.harvard.edu/barro/publications/religion-adherence-data.

Abortion. The abortion index (Bloom et al., 2009) measures the legal availability of abortion. The index takes integer values in the range 0-7. For each of seven reasons each country receives a score of 1 if abortion is allowed, while if it is not permitted the score is 0. The index is constructed as the summation of the scores for the seven reasons. The reasons are: to 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). 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). Dataset available at: https://www.hsph.harvard.edu/david-canning/data-sets/

Women vote. The year in which women acquired the rights to vote and to stand for election. Source: Inter-Parliamentary Union, https://www.ipu.org/.

Predicted mortality. Acemoglu and Johnson (2007) construct the index predicted mortality to capture the dramatic reduction in the mortality from infectious diseases occurred in the world in the 1940s and 1950s. Predicted mortality 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: $\mathrm{M}_{it}^{I} = \sum_{d \in D} \left[(1 - I_{di}) \, M_{di40} + I_{di} M_{dFt} \right]$, where M_{di40} denotes mortality in country i from disease d at year 1940, M_{dFt} is the mortality rate from disease d at the health frontier of the world at time t, I_{di} is a dummy taking the value zero prior to the global intervention for disease d and a value of one after the intervention, and D denotes the set of 15 infectious diseases considered in the paper. The dataset of Acemoglu and Johnson (2007) presents time series observations on predicted mortality at 10-year intervals from 1940 to 2000, and has been downloaded from: https://economics.mit.edu/faculty/acemoglu/data.

World income. This variable is the *Trade-Weighted World Income Instrument* (Acemoglu et al., 2008) constructed for country i and 5 year period t as: $\sum_{i,j \neq i}^{N} \omega_{i,j} Y_{j,t}$, were $\omega_{i,j}$ is the share of trade between country i and country j in the GDP of country i, and $Y_{j,t}$ denotes log total income. Each observation is taken every fifth year from 1950 to 2000. Data were downloaded from: https://economics.mit.edu/faculty/acemoglu/data.

Oil shock. The change in the international oil price between year t and $t - 5$ multiplied by countries' average GDP share of net oil exports (Brueckner and Schwandt, 2015). Oil shock data were downloaded from The Economic Journal's website.

School_40-64. The total average years of schooling of men and women aged 40-64. Calculated from data on educational attainment and population disaggregated by sex and by 5-year age intervals. Source: Barro and Lee (2013), http://www.barrolee.com/.

Export. The share in current GDP of merchandise export at current PPP. Source: Penn World Table version 9.0, variable csh_x, (Feenstra, Inklaar and Timmer, 2015), https://www.rug.nl/ggdc/productivity/pwt/. In Lewbel's regressions, each cross-sectional observation is the average of the annual data between 1956 and 1999.

C2 Robustness to outliers

We investigate the robustness to outliers of the estimates of our baseline model (column 7 in Table 8) with the GMM estimation of the equation on 36 samples obtained by deleting one country in each regression (see Young, 2018 for a deep investigation of the negative effects of non-iid errors on the quality of inference in 2SLS estimates). Table C2 presents a summary of the effect of deleting one country on both the p-values of the three endogenous regressors (Panel a) and the F-statistic proposed by Sanderson and Windmeijer (2016) to test the weak IV hypothesis in first stage equations (Panel b). The statistical significance of the coefficients of the three main variables is confirmed in all but one case that gives a p-value of the coefficient of log infant mortality of 0.105. From the first stage regressions of Log(GDP) and Log(infant mortality), we get F-statistics greater than 10 in two-thirds of the samples. In the same regressions, another 23% of the samples give Fs in the range 9 to 10. The first stage equation of schooling shows no evidence of weak IV in all the 36 samples. We also test for the normality of the residuals of the GMM estimation of the baseline equation on the full sample by applying the test proposed by D'Agostino, Belanger, and D'Agostino (1990) and implemented by the STATA command sktest. The test-statistic has approximately a χ^2 distribution with 2 degrees of freedom under the null of normality, and in the full sample estimates, it takes the value of 1.86 with a p-value of 0.39.

REFERENCES

D'Agostino, Ralph B., Albert Belanger, and Ralph B. D'Agostino Jr., "A suggestion for using powerful and informative tests of normality," *The American Statistician* 44:4 (1990), 316-321.

Feenstra, Robert C., Robert Inklaar, and Marcel P. Timmer, "The next generation of the Penn World Table," American economic review 105:10 (2015), 3150-3182.

Perron Pierre and Xiaokang Zhu, "Structural Breaks with Deterministic and Stochastic Trends," *Journal of Econometrics* 129 (2005) 65-119.

Perron Pierre and Tomoyoshi Yabu, "Testing for Shifts in Trend with an Integrated or Stationary Noise Component," *Journal of Business and Economic Statistics* 27 (2009), 369-396.

Roy, Anindya and Wayne A., Fuller, "Estimation for Autoregressive Processes With a Root Near One," *Journal of Business & Economic Statistics* 19:4 (2001), 482–493.

Wang, Qunyong, and Na Wu, "Long-run covariance and its applications in cointegration regression," *The Stata Journal* 12:3 (2012), 515-542.

Table A1. List of countries in convergence analysis

Afghanistan	Canada	Gabon	Lao PDR	Nicaragua	Sri Lanka
Albania	Central African Republic	Gambia, The	Latvia	Niger	Sudan
Algeria	Chad	Georgia	Lebanon	Nigeria	Suriname
Angola	Channel Islands	Germany	Lesotho	Norway	Swaziland
Argentina	Chile	Ghana	Liberia	Oman	Sweden
Armenia	China	Greece	Libya	Pakistan	Switzerland
Aruba	Colombia	Grenada	Lithuania	Panama	Syrian Arab Republic
Australia	Comoros	Guam	Luxembourg	Papua New Guinea	Tajikistan
Austria	Congo, Dem. Rep.	Guatemala	Macedonia, FYR	Paraguay	Tanzania
Azerbaijan	Congo, Rep.	Guinea	Madagascar	Peru	Thailand
Bahamas, The	Costa Rica	Guinea-Bissau	Malawi	Philippines	Timor-Leste
Bahrain	Cote d'Ivoire	Guyana	Malaysia	Poland	Togo
Bangladesh	Croatia	Haiti	Maldives	Portugal	Tonga
Barbados	Cuba	Honduras	Mali	Puerto Rico	Trinidad and Tobago
Belarus	Cyprus	Hungary	Malta	Qatar	Tunisia
Belgium	Czech Republic	Iceland	Mauritania	Romania	Turkey
Belize	Denmark	India	Mauritius	Russian Federation	Turkmenistan
Benin	Djibouti	Indonesia	Mexico	Rwanda	Uganda
Bhutan	Dominican Republic	Iran, Islamic Rep.	Micronesia, Fed. Sts.	Samoa	Ukraine
Bolivia	Ecuador	Iraq	Moldova	Sao Tome and Principe	United Arab Emirates
Bosnia and Herzegovina	Egypt, Arab Rep.	Ireland	Mongolia	Saudi Arabia	United Kingdom
Botswana	El Salvador	Israel	Montenegro	Senegal	United States of America
Brazil	Equatorial Guinea	Italy	Morocco	Sierra Leone	Uruguay
Brunei Darussalam	Eritrea	Jamaica	Mozambique	Singapore	Uzbekistan
Bulgaria	Estonia	Japan	Myanmar	Slovak Republic	Vanuatu
Burkina Faso	Ethiopia	Jordan	Namibia	Slovenia	Venezuela, RB
Burundi	Fiji	Kenya	Nepal	Solomon Islands	Vietnam
Cabo Verde	Finland	Korea, Dem. People’s Rep.	Netherlands	Somalia	Yemen, Rep.
Cambodia	France	Korea, Rep.	New Caledonia	South Africa	Zambia
Cameroon	French Polynesia	Kuwait	New Zealand	Spain	Zimbabwe

Table A2. List of countries in convergence groups

Group H	Group I		Group L	
		Group H-L		Group L-L
Niger	Tanzania	Samoa	Sri Lanka	Gabon
Somalia	Afghanistan	Ghana	Turkey	Israel
Congo, Dem. Rep.	Benin	Kenya	Bahrain	Argentina
Mali	Zambia	Solomon Islands	Malaysia	New Caledonia
Chad	Cote d'Ivoire	Zimbabwe	French Polynesia	Jamaica
Burundi	Guinea	Papua New Guinea	Kuwait	Uruguay
Angola	Central African Republic	Tonga	Qatar	New Zealand
Uganda	Senegal	Pakistan	Vietnam	France
Timor-Leste	Cameroon	Jordan	Colombia	Georgia
Nigeria	Malawi	Namibia	Brunei Darussalam	Azerbaijan
Burkina Faso	Guinea-Bissau	Tajikistan	Costa Rica	Ireland
Gambia, The	Mauritania	Vanuatu	United Arab Emirates	Iceland
Mozambique	Congo, Rep.	Egypt, Arab Rep.	Aruba	Korea, Dem. People's Rep.
Equatorial Guinea	Sierra Leone	Micronesia, Fed. Sts.	Trinidad and Tobago	Sweden
	Liberia	Swaziland	Brazil	Australia
	Sudan	Lesotho	Iran, Islamic Rep.	United States
	Togo	Guatemala	Albania	United Kingdom
	Sao Tome and Principe	Haiti	Lebanon	Chile
	Comoros	Syrian Arab Republic	China	Bahamas, The
	Ethiopia	Djibouti	Thailand	Norway
	Iraq	Philippines	Mauritius	Barbados
	Madagascar	Bolivia	Korea, Rep.	Belgium
	Eritrea	Turkmenistan	Singapore	Finland
	Yemen, Rep.	Algeria		Netherlands
	Rwanda	Lao PDR		Montenegro
		Botswana		Denmark
		Oman		Cuba
		Mongolia		Russian Federation
		Saudi Arabia		Armenia
		Cambodia		Belarus
		Honduras		Canada
		Belize		Lithuania
		Fiji		Slovenia
		Paraguay		Estonia
		Guyana		Switzerland
		Panama		Luxembourg
		Ecuador		Latvia
		Morocco		Bulgaria
		South Africa		Ukraine
		Dominican Republic		Macedonia, FYR
		Peru		Puerto Rico
		Cabo Verde		Romania
		Suriname		Czech Republic
		India		Croatia
		Indonesia		Channel Islands
		Guam		Austria
		Venezuela, RB		Germany
		Libya		Japan
		Uzbekistan		Malta
		Nepal		Italy
		Nicaragua		Cyprus
		Myanmar		Slovak Republic
		Mexico		Hungary
		Tunisia		Greece
		Bangladesh		Bosnia and Herzegovina
		Bhutan		Poland
		Maldives		Spain
		Grenada		Moldova
		El Salvador		Portugal

In each group, countries are listed in decreasing order of the last five years' average TFR.

Table B1. Onset year of fertility transition

Country	Onset year	W_{RQF}	β_1	β_2	Country	Onset year	W_{RQF}	β_1	β_2
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 Repub.	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, Rep.	1972	59.449	0.006	-0.007					
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	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
Djibouti	1987	66.272	0.001	-0.030	Sri Lanka	1956	1282.938	-0.007	-0.019
Philippines	1961	66.985	-0.004	-0.017	Turkey	1966	62.897	-0.008	-0.024
Bolivia	1982	48.874	-0.007	-0.020	Bahrain	1966	76.988	0.002	-0.027
Turkmenistan	1964	91.099	0.026	-0.022	Malaysia	1960	76.074	0.001	-0.021
Algeria	1973	57.451	0.003	-0.031	French Polynesia	1969	59.170	-0.009	-0.022
Lao PDR	1990	75.959	0.002	-0.035	Kuwait	1967	4.200	0.001	-0.030
Botswana	1976	65.419	0.000	-0.025	Qatar	1975	72.284	-0.002	-0.032
Oman	1985	75.216	0.004	-0.043	Vietnam	1969	86.241	0.014	-0.033
Mongolia	1968	69.779	0.023	-0.031	Colombia	1960	77.375	0.002	-0.026
Saudi Arabia	1983	72.654	0.000	-0.033	Brunei Darussalam	1963	68.954	-0.002	-0.028
Cambodia	1988	57.703	-0.005	-0.032	Costa Rica	1957	1302.508	0.020	-0.023
Honduras	1979	61.695	-0.004	-0.027	United Arab Emirates	1981	62.152	-0.007	-0.039
Belize	1977	65.000	-0.003	-0.024	Trinidad and Tobago	1956	3158.447	-0.002	-0.022
Fiji	1956	35.760	-0.013	-0.016	Brazil	1963	73.992	-0.001	-0.026
Paraguay	1988	44.784	-0.010	-0.025	Iran, Islamic Rep.	1983	73.015	-0.004	-0.049
Guyana	1956	456524.906	0.019	-0.018	Albania	1958	41533.449	0.015	-0.027
Panama	1958	320.260	0.006	-0.017	Lebanon	1968	63.088	-0.004	-0.028
Ecuador	1964	69.737	-0.001	-0.021	China	1963	0.000	-0.006	-0.031
Morocco	1969	69.742	0.003	-0.026	Thailand	1962	74.425	0.002	-0.032
South Africa	1969	63.565	-0.002	-0.021	Mauritius	1956	14.826	-0.006	-0.026
Dominican Republic	1958	1075.737	0.001	-0.022	Korea, Rep.	1956	15.084	0.034	-0.034

In each group, countries are listed in decreasing order of the last five years' average TFR. W_{RQF} is the Wald robust quasi-FGLS statistic proposed by Perron and Yabu (2009) to test the null that the trend function does not change in the period under consideration. The test is asymptotically distributed as a chi-squared random variable. Critical values: 2.706 at 10%, 3.841 at 5%, 6.635 at 1%.

Table C1. List of countries in regressions of fertility transition onsets

Albania	Indonesia	Peru
Algeria	Iran, Islamic Republic	Philippines
Bahrain	Iraq	Qatar
Bangladesh	Jordan	Rwanda
Benin	Kenya	Saudi Arabia
Bolivia	Korea, Republic	Senegal
Botswana	Kuwait	Sierra Leone
Brazil	Lao PDR	Singapore
Cambodia	Lesotho	South Africa
Cameroon	Liberia	Sri Lanka
Central African Republic	Libya	Sudan
China	Malawi	Swaziland
Colombia	Malaysia	Syrian Arab Republic
Congo, Rep.	Mauritania	Tanzania
Costa Rica	Mauritius	Thailand
Cote d'Ivoire	Mexico	Togo
Dominican Republic	Mongolia	Trinidad and Tobago
Ecuador	Morocco	Tunisia
Egypt, Arab Republic	Myanmar	Turkey
El Salvador	Namibia	United Arab Emirates
Ghana	Nepal	Venezuela, RB
Guatemala	Nicaragua	Vietnam
Haiti	Pakistan	Yemen, Rep.
Honduras	Panama	Zambia
India	Paraguay	Zimbabwe

Table C2. Determinants of fertility transition onsets. First stage in 2SLS panel within estimates. Dependent variables: Log(*GDP*), Log(infant mortality), Log(*s*chool_15-39).

	(1)			(2)			(3)		
Equation	GDP p. c.	Infant mortality	Schooling	GDP p. c.	Infant mortality	Schooling	GDP p. c.	Infant mortality	Schooling
Log(Urbanization)				0.253 (0.320)	-0.245 (0.213)	0.284 (0.231)	0.150 (0.274)	-0.205 (0.195)	0.325 (0.213)
Log(Abortion)							0.087* (0.046)	-0.022 (0.036)	-0.027 (0.046)
Log(*School_40-64)*	0.247* (0.136)	-0.372*** (0.088)	0.560*** (0.146)	0.190 (0.117)	-0.316*** (0.078)	0.496*** (0.140)	0.194 (0.130)	-0.332*** (0.069)	0.477*** (0.145)
Log(*School_40-64)**D1960-1964	-0.049 (0.032)	0.129*** (0.031)	-0.029 (0.023)	-0.040 (0.029)	0.121*** (0.025)	-0.019 (0.019)	-0.029 (0.034)	0.119*** (0.021)	-0.026 (0.021)
Log(*School_40-64)**D1970-1974	-0.041 (0.052)	0.003 (0.040)	0.139*** (0.042)	-0.036 (0.048)	-0.001 (0.038)	0.144*** (0.042)	-0.019 (0.046)	-0.013 (0.037)	0.142*** (0.042)
World income	0.328*** (0.036)	-0.273*** (0.036)	-0.047 (0.040)	0.351*** (0.047)	-0.296*** (0.041)	-0.022 (0.041)	0.316*** (0.035)	-0.271*** (0.029)	-0.012 (0.037)
(World income)*South Asia&East Asia&Pacific	0.352* (0.183)	-0.149 (0.100)	0.087 (0.087)	0.304 (0.209)	-0.102 (0.118)	0.033 (0.114)	0.380* (0.200)	-0.126 (0.104)	0.007 (0.106)
(World income)*D1970-1974	-0.002 (0.002)	0.002** (0.001)	-0.000 (0.001)	-0.002 (0.002)	0.002** (0.001)	-0.000 (0.001)	-0.003* (0.002)	0.003** (0.001)	0.000 (0.001)
Predicted mortality	-2.888*** (0.831)	-2.376*** (0.677)	0.959 (1.297)	-2.797*** (0.898)	-2.464*** (0.835)	1.062 (0.982)	-3.053*** (0.887)	-2.508*** (0.843)	1.147 (0.969)
Predicted mortality*Latin America	0.122 (0.112)	0.237** (0.100)	0.093 (0.124)	0.127 (0.124)	0.232** (0.101)	0.099 (0.114)	0.181 (0.123)	0.232*** (0.082)	0.039 (0.121)
Predicted mortality *Middle East& North Africa	-0.132 (0.284)	0.621*** (0.224)	-1.512*** (0.419)	-0.144 (0.265)	0.632*** (0.201)	-1.525*** (0.413)	-0.115 (0.235)	0.615*** (0.192)	-1.554*** (0.405)
Predicted mortality* Sub-Saharan Africa	-0.113 (0.108)	0.332*** (0.066)	0.182* (0.108)	-0.119 (0.107)	0.338*** (0.063)	0.175* (0.091)	-0.175 (0.118)	0.356*** (0.061)	0.200** (0.098)

Table C2 continued

Predicted mortality *Europe&Central Asia	-0.361***	0.451***	-0.474***	-0.302	0.394***	-0.408***	-0.399**	0.418***	-0.375**
	(0.126)	(0.088)	(0.133)	(0.197)	(0.131)	(0.155)	(0.175)	(0.117)	(0.154)
Predicted mortality*D1960-1964	2.963***	2.156***	-1.362	2.898***	2.219***	-1.436	3.109***	2.282***	-1.515*
	(0.772)	(0.655)	(1.200)	(0.825)	(0.797)	(0.910)	(0.817)	(0.808)	(0.894)
Predicted mortality*D1965-1969	2.905***	2.253***	-1.213	2.834***	2.322***	-1.293	3.045***	2.377***	-1.364
	(0.782)	(0.654)	(1.210)	(0.832)	(0.801)	(0.913)	(0.830)	(0.813)	(0.895)
Sanderson-Windmeijer test F-statistic	10.690	10.502	172.772	6.175	8.567	73.996	6.776	10.545	66.924

Instruments: world income plus two variables obtained from the interactions between world income and the dummies for the period 1970--1974 and for the countries of South Asia, East Asia and Pacific; 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; log school_40--64, and its product with the dummies for the period 1960-1964 and 1970-1974.
All regressions include time dummies. Standard errors are robust to heteroskedasticity and clustered at the country level. *t* statistics in parentheses. S-W is the Sanderson-Windmeijer test for weak instruments.
Standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table C3. Second stage coefficient p-value and first stage F-statistic from GMM estimates of the baseline model with the removal of one country from the full sample

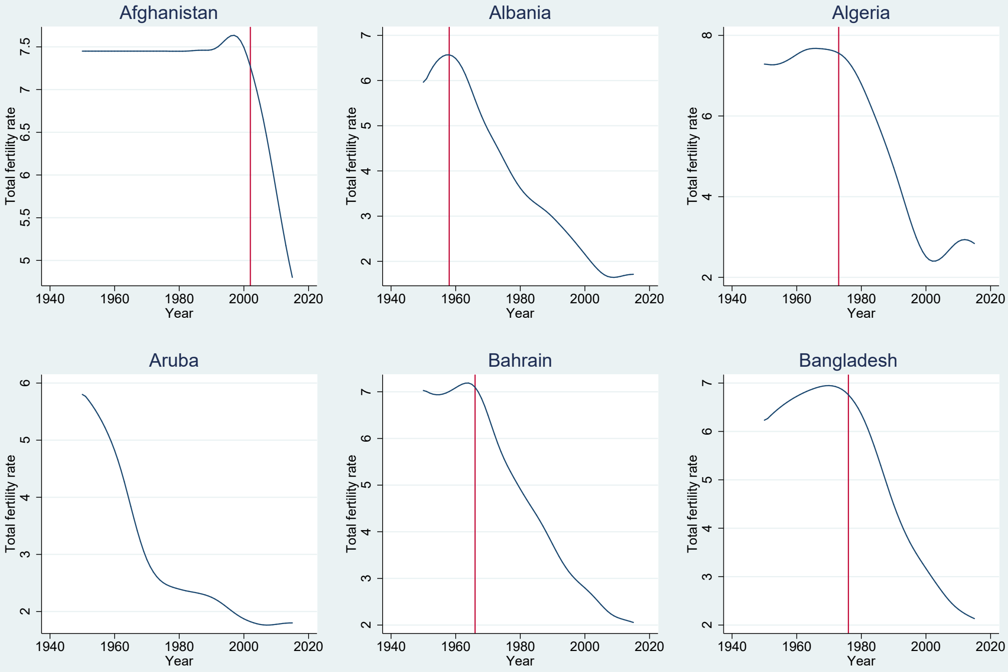
Panel A			
Variable in second-stage equation	Log(GDP per capita)	Log infant mortality	Log (schoolw_15-39)
Maximum p-value	0.004	0.105	0
Minimum p-value	0	0.002	0
Panel B			
First-stage equation	Log(GDP per capita)	Log infant mortality	Log (schoolw_15-39)
Sanderson-Windmeijer test F-statistic			
Maximum	16.04	18.49	17206.85
Minimum	5.41	6.47	52.94
Regressions* with			
F < 8	0.03	0.06	0
8 < F < 9	0.09	0.06	0
9 < F < 10	0.23	0.23	0
F > 10	0.66	0.66	1

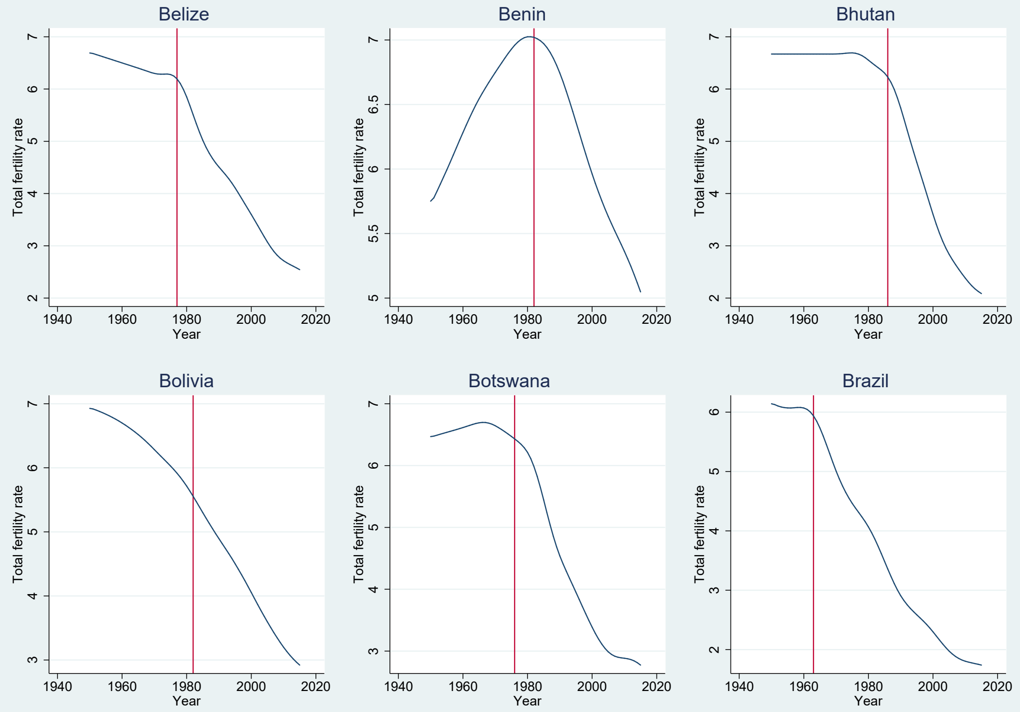
*Relative frequency

Table C4. Cross-country correlation matrix

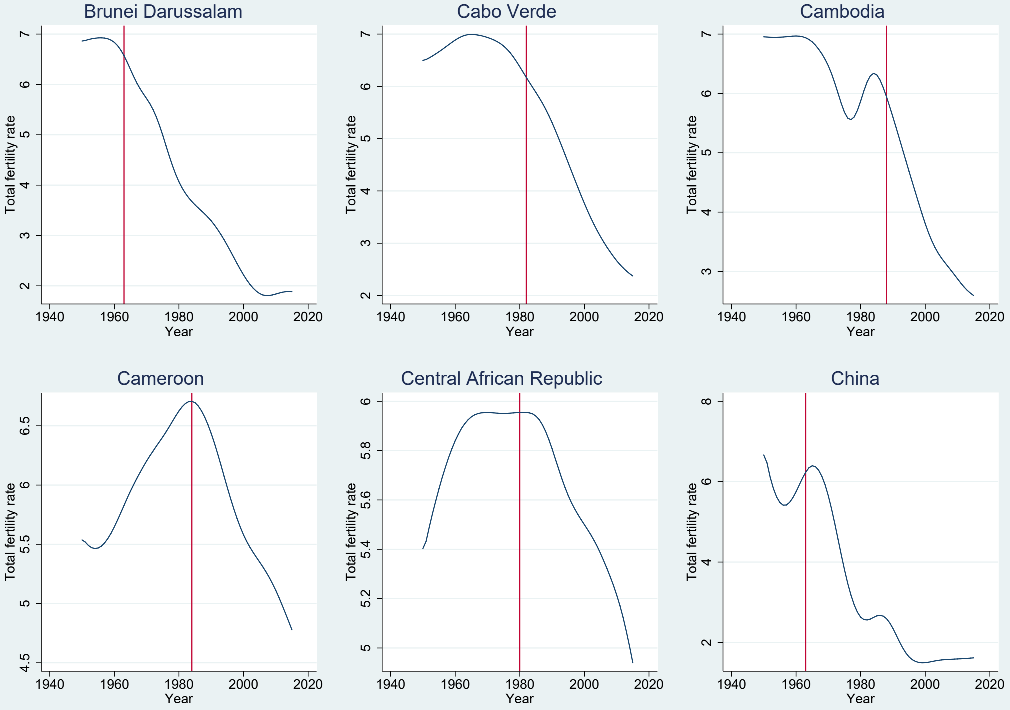
	Countries	Transition	GDP p. c.	Infant mortality	Schoolw _15-39	Urbanization	Catholic	Muslim	Women vote	Abortion
Transition	75	1								
GDP per capita	75	0.206	1							
Infant mortality	75	-0.723	-0.526	1						
Schoolw_15-39	75	0.661	0.191	-0.708	1					
Urbanization	75	0.377	0.700	-0.567	0.193	1				
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

Figure B1. Fertility and Onset Year of Fertility Decline in Transitioning Countries

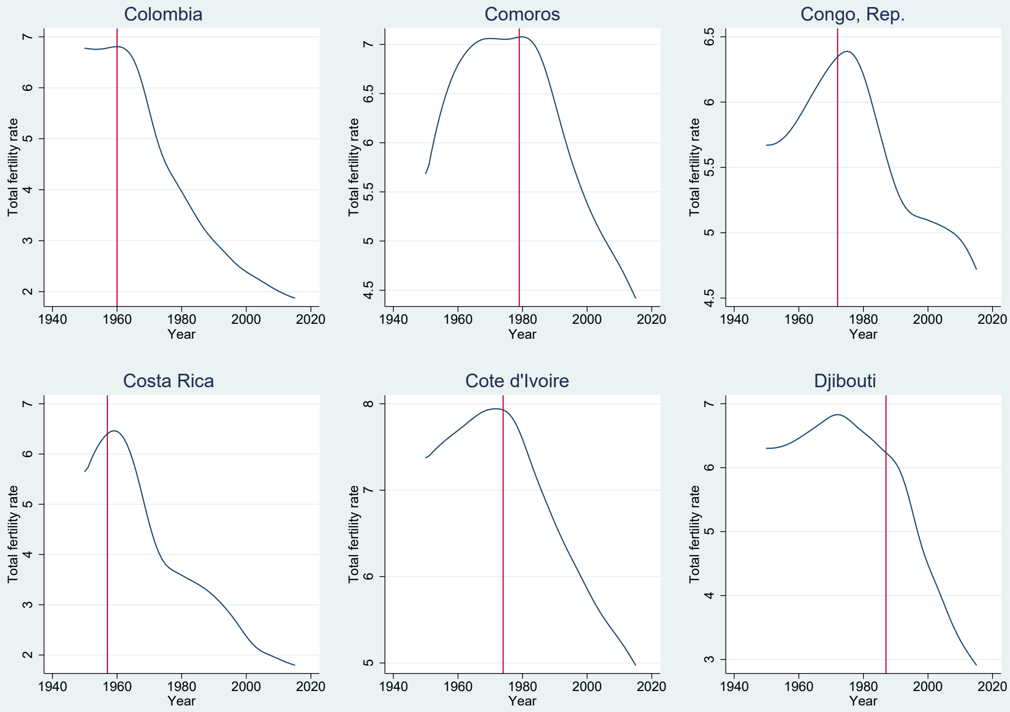


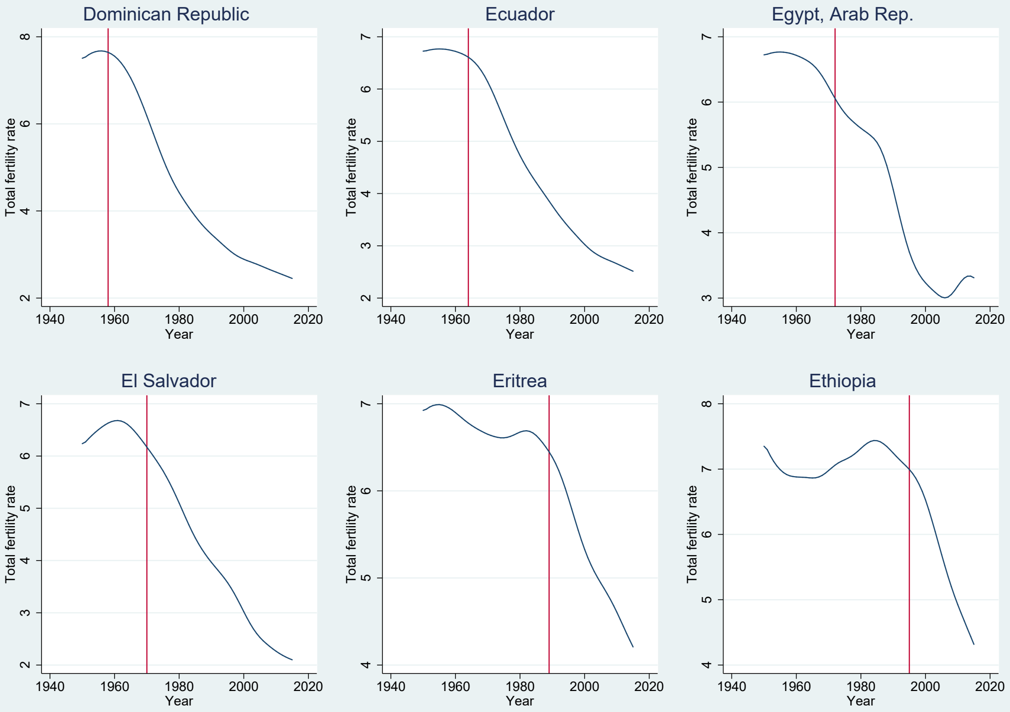


Belize
Benin
Bhutan
Bolivia
Botswana
Brazil
Total fertility rate
Year
1940
1960
1980
2000
2020

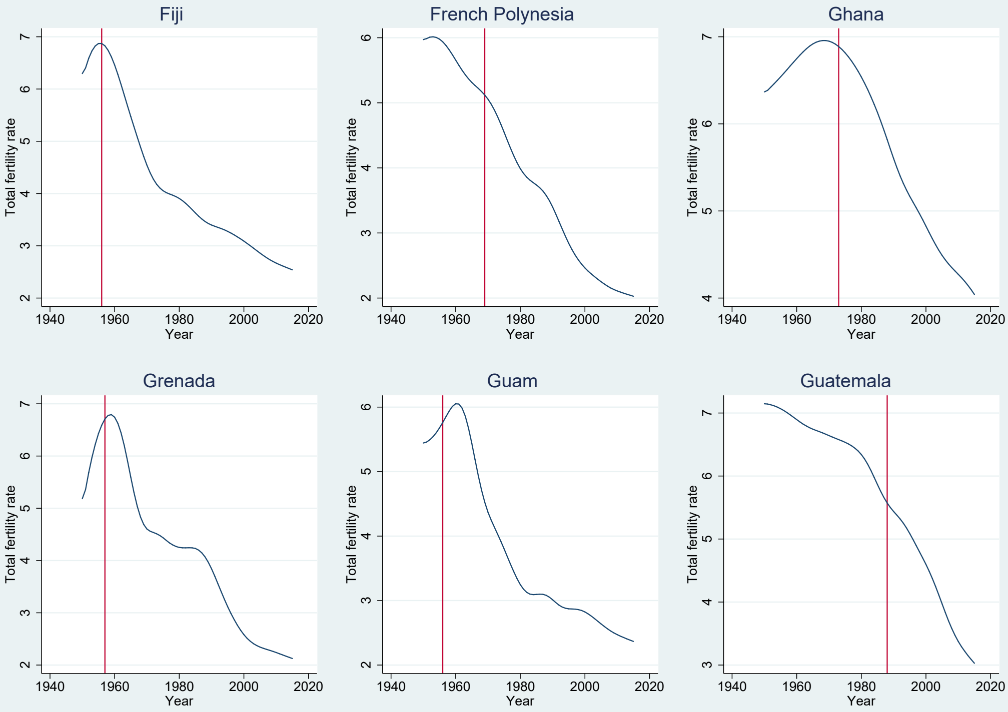


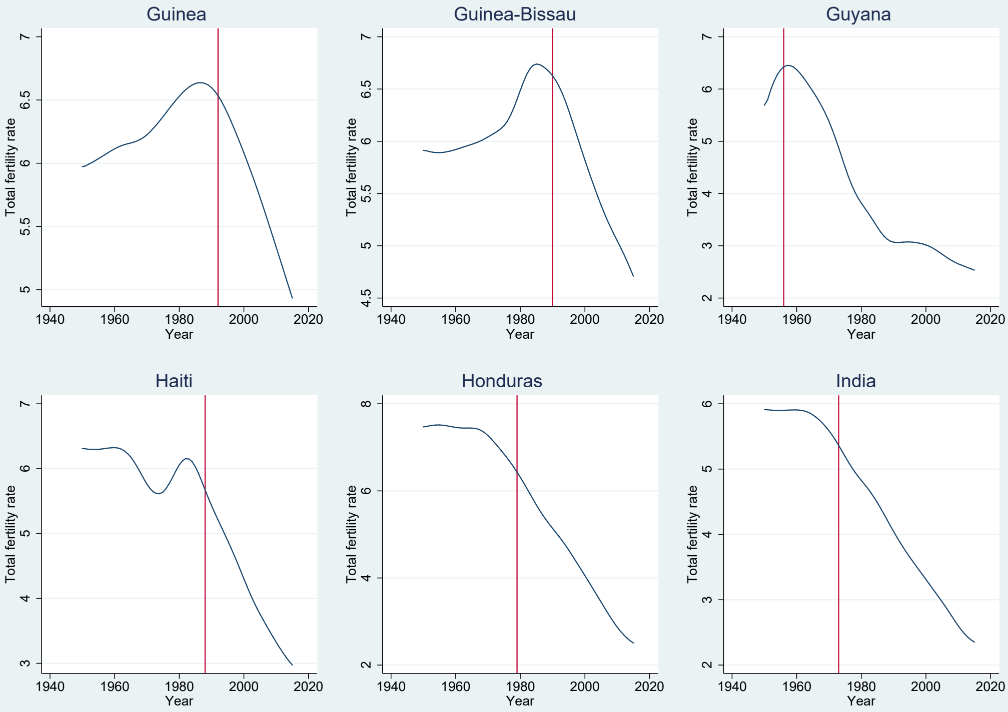
Brunei Darussalam
Cabo Verde
Cambodia
Cameroon
Central African Republic
China
Total fertility rate
Year
1940
1960
1980
2000
2020


Colombia
Comoros
Congo, Rep.
Costa Rica
Cote d'Ivoire
Djibouti
Total fertility rate
Year
1940
1960
1980
2000
2020

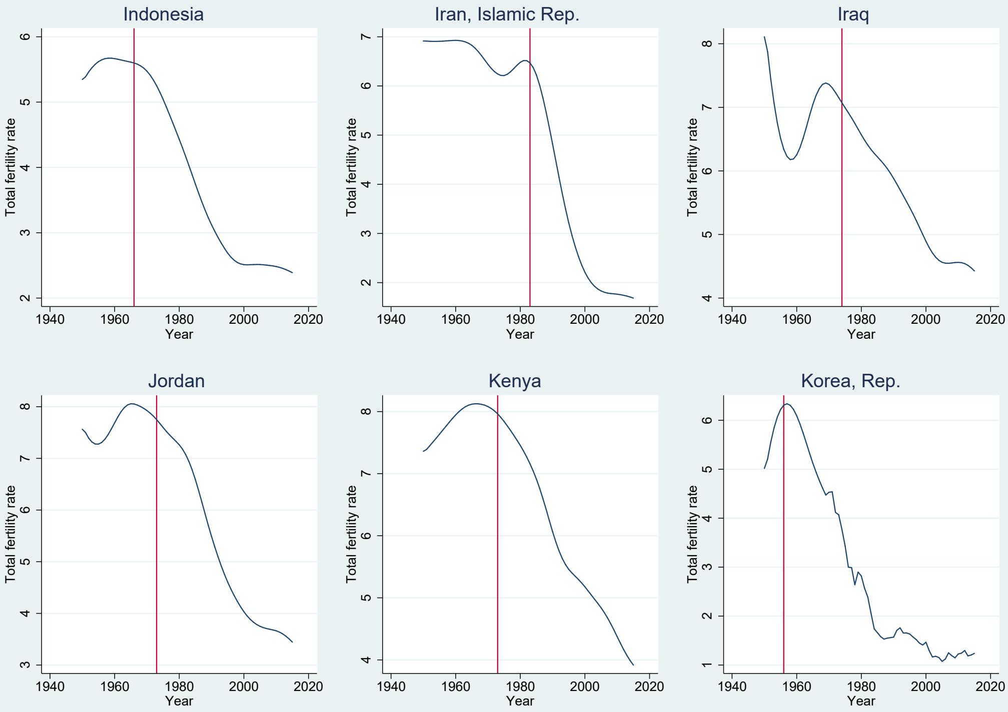


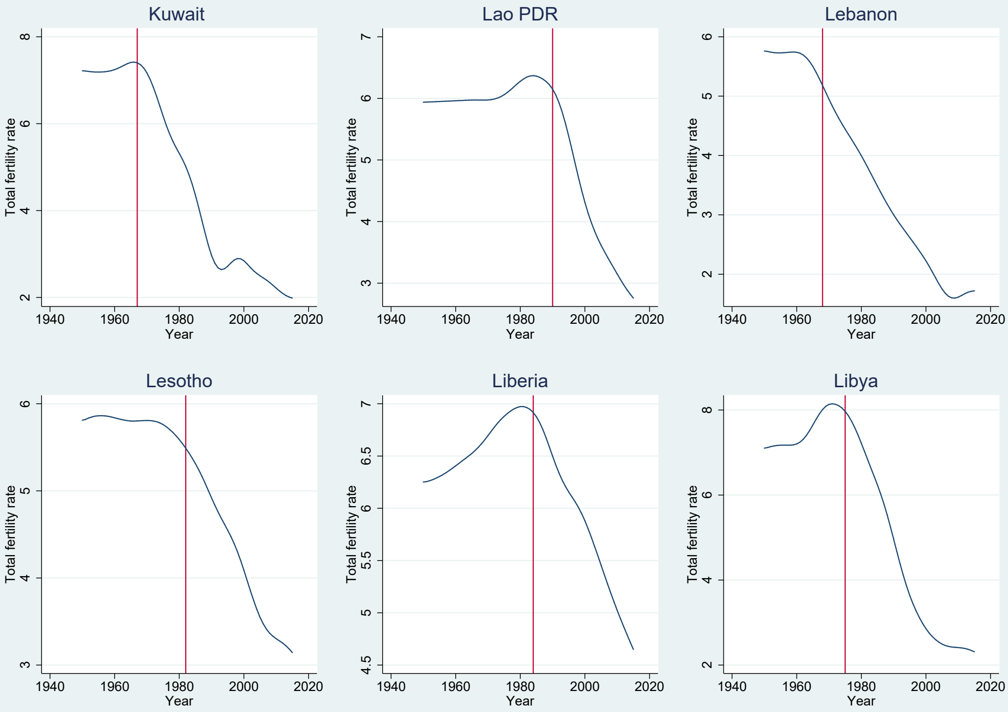
Dominican Republic
Ecuador
Egypt, Arab Rep.
El Salvador
Eritrea
Ethiopia
Total fertility rate
Year
1940
1960
1980
2000
2020


Fiji
French Polynesia
Ghana
Grenada
Guam
Guatemala
Total fertility rate
Year
1940
1960
1980
2000
2020

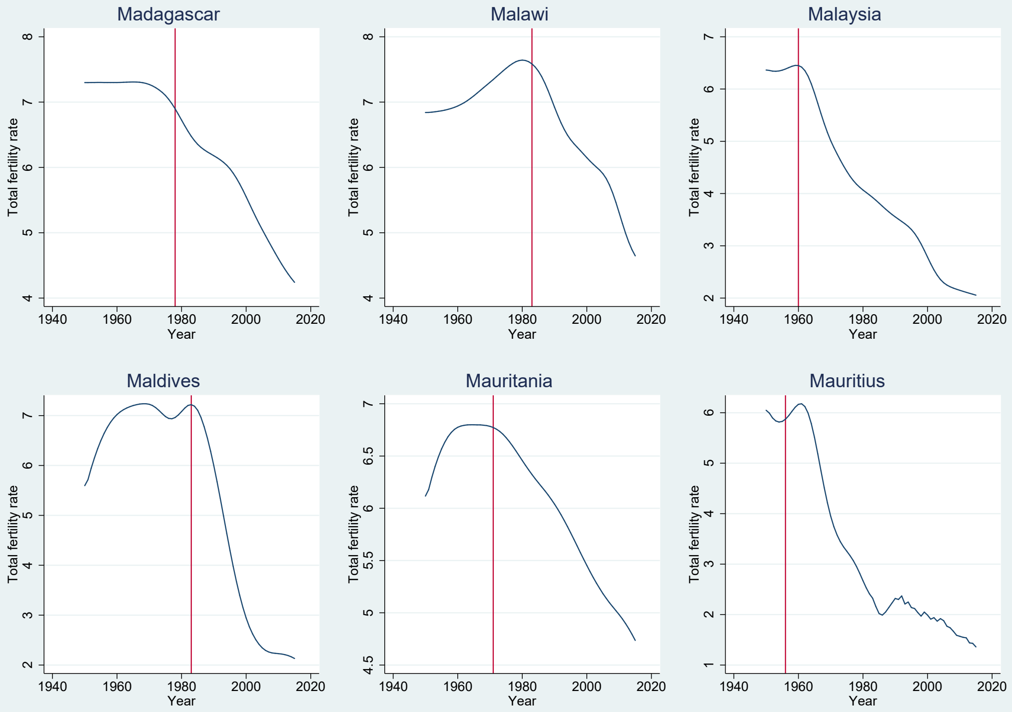


Guinea
Guinea-Bissau
Guyana
Haiti
Honduras
India
Total fertility rate
Year
1940
1960
1980
2000
2020

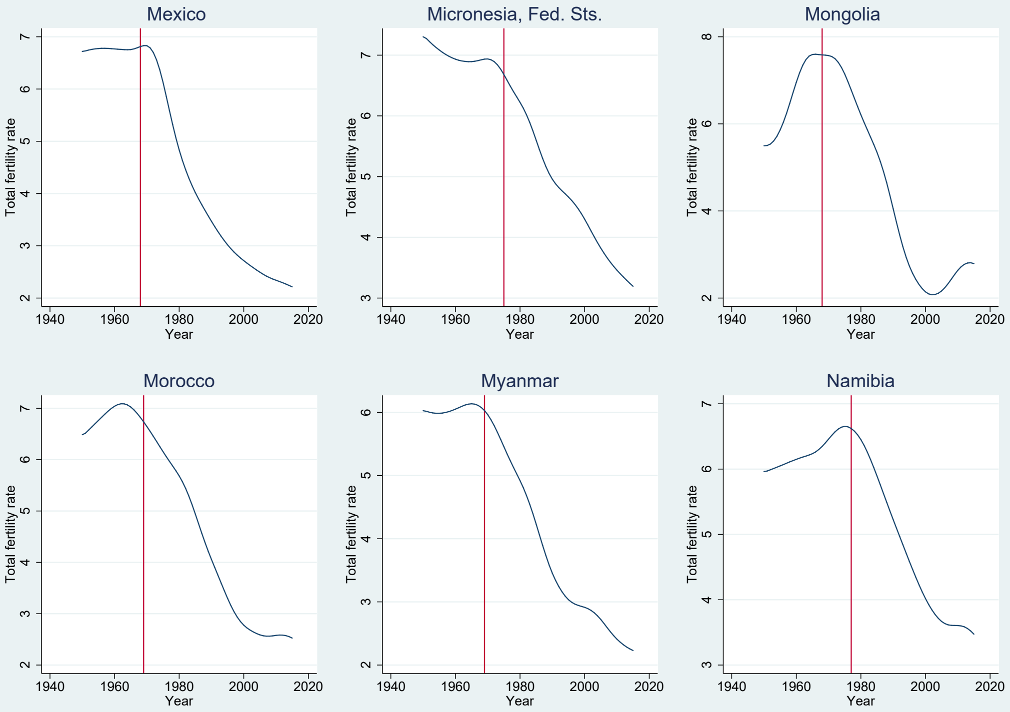

Indonesia
Iran, Islamic Rep.
Iraq
Jordan
Kenya
Korea, Rep.
Total fertility rate
Year
1940
1960
1980
2000
2020

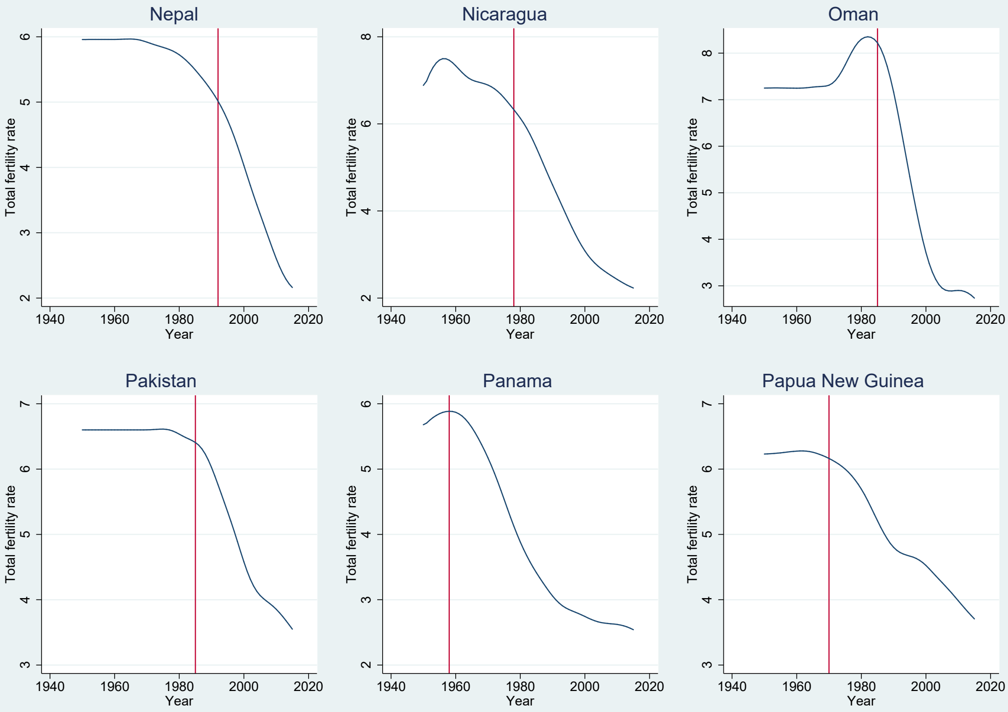


Kuwait
Lao PDR
Lebanon
Lesotho
Liberia
Libya
Total fertility rate
Year
1940
1960
1980
2000
2020

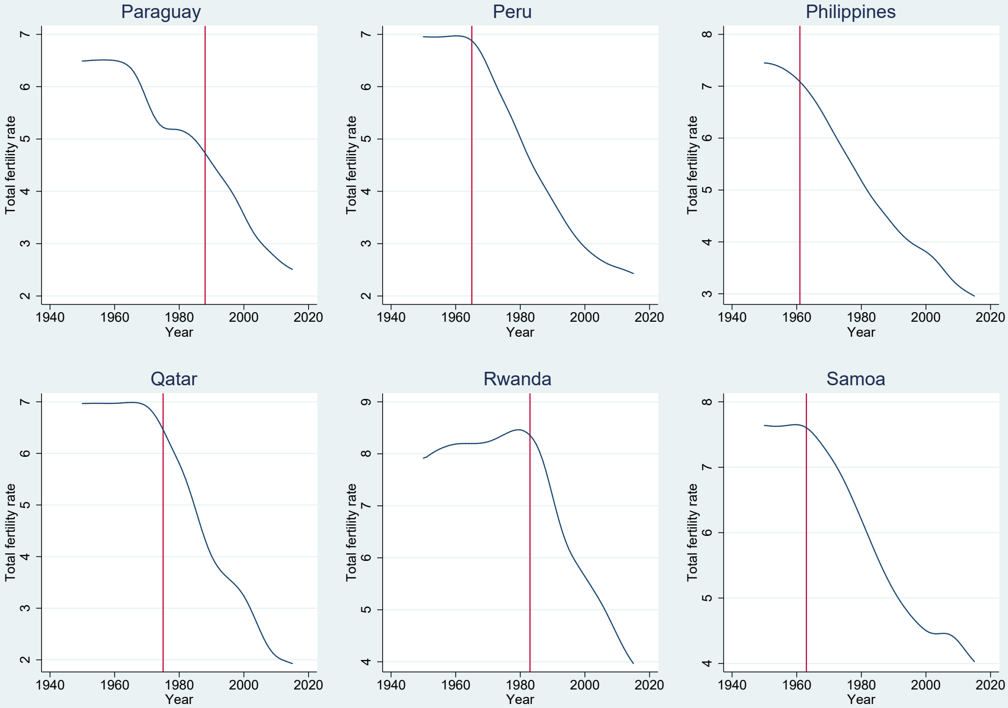


Madagascar
Malawi
Malaysia
Maldives
Mauritania
Mauritius
Total fertility rate
Year
1940
1960
1980
2000
2020

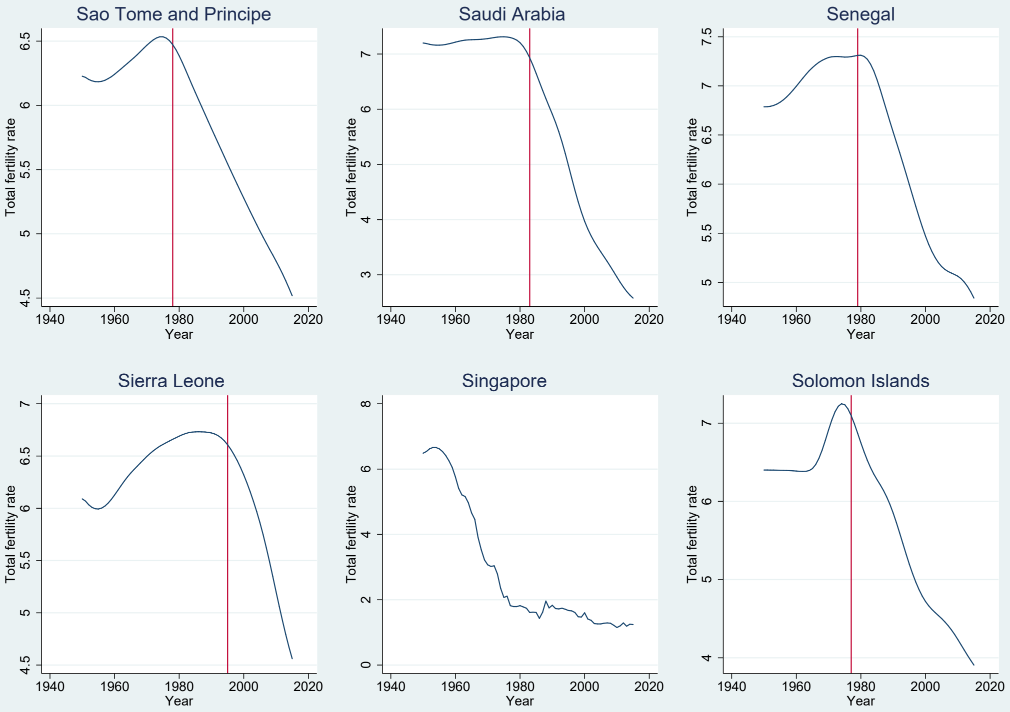

Mexico
Micronesia, Fed. Sts.
Mongolia
Morocco
Myanmar
Namibia
Total fertility rate
Year
1940
1960
1980
2000
2020



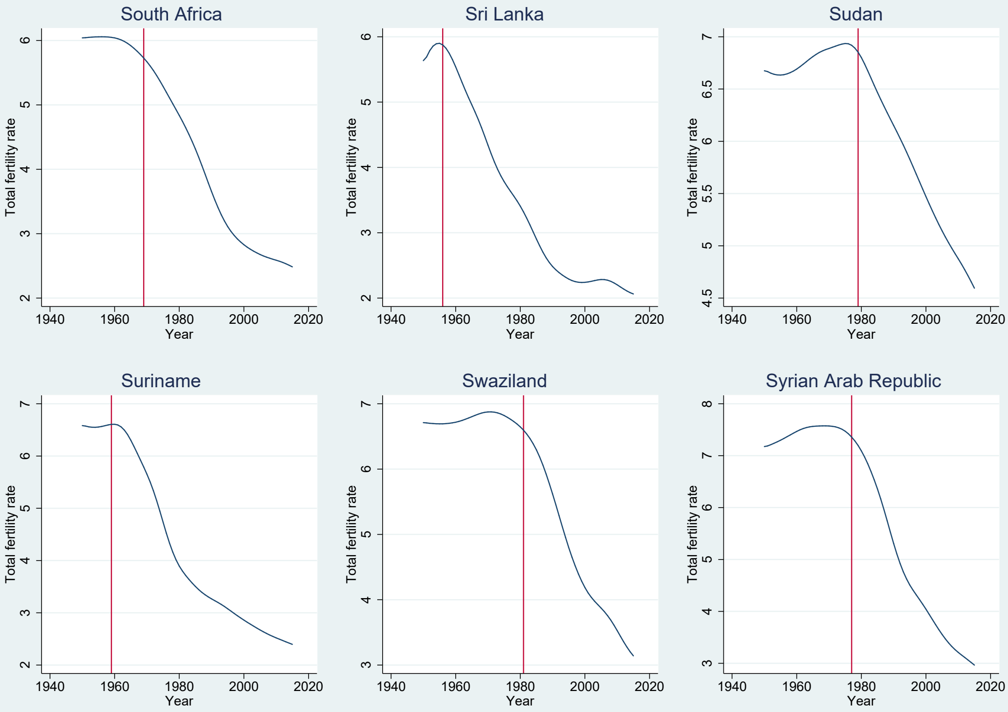
Nepal
Nicaragua
Oman
Pakistan
Panama
Papua New Guinea
Total fertility rate
Year
1940
1960
1980
2000
2020

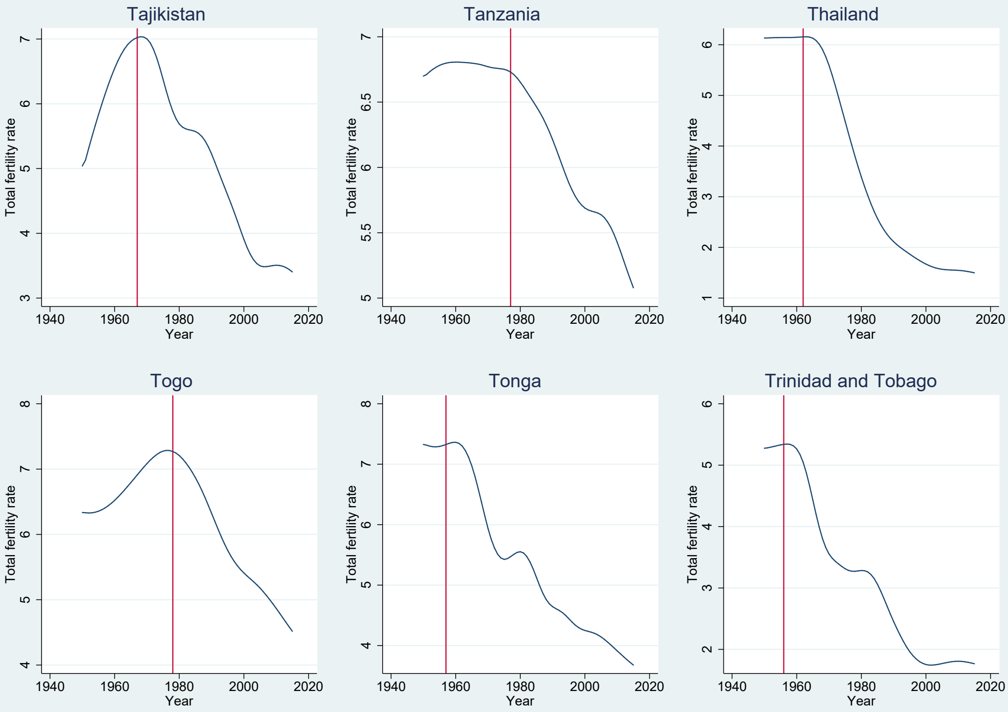


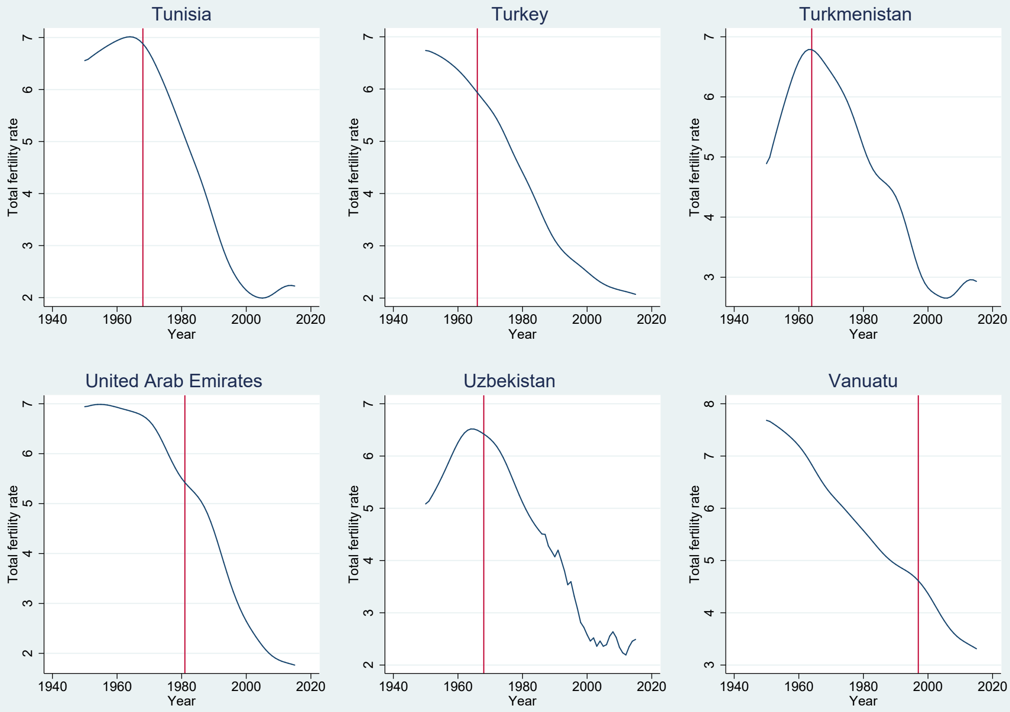
Paraguay
Peru
Philippines
Qatar
Rwanda
Samoa
Total fertility rate
Year
1940
1960
1980
2000
2020
2
3
4
5
6
7
8
9

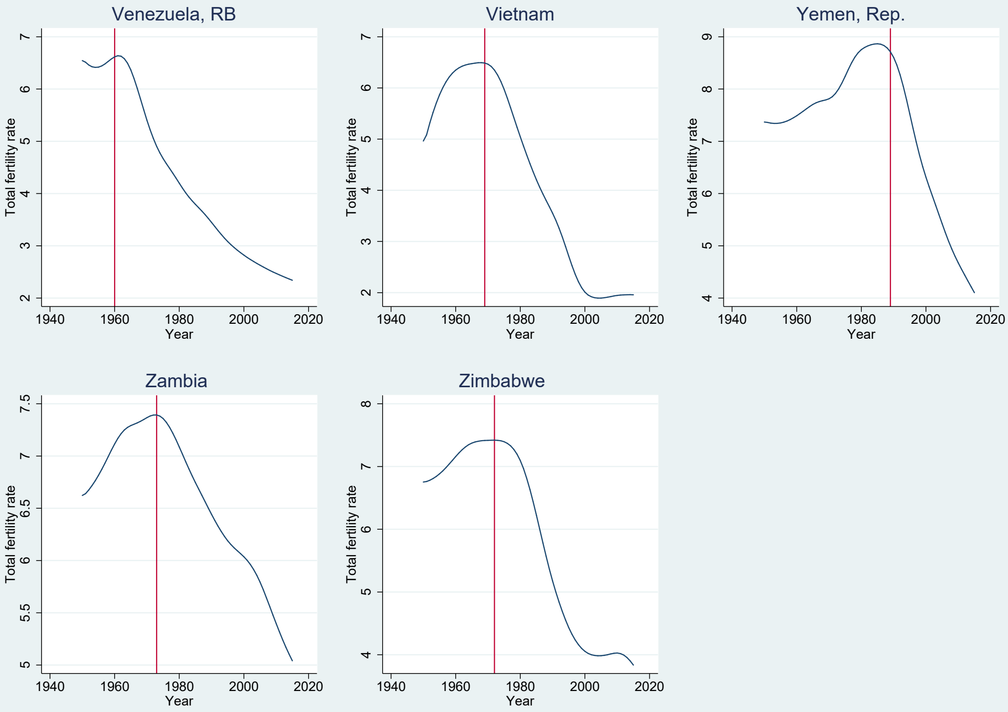


Sao Tome and Principe
Saudi Arabia
Senegal
Sierra Leone
Singapore
Solomon Islands
Total fertility rate
Year
1940
1960
1980
2000
2020


South Africa
Sri Lanka
Sudan
Suriname
Swaziland
Syrian Arab Republic
Total fertility rate
Year
1940
1960
1980
2000
2020


Tajikistan
Tanzania
Thailand
Togo
Tonga
Trinidad and Tobago
Total fertility rate
Year
1940
1960
1980
2000
2020
3
4
5
6
7
6.5
5.5
1
2
8


Tunisia
Turkey
Turkmenistan
United Arab Emirates
Uzbekistan
Vanuatu
Total fertility rate
Year


Venezuela, RB
Vietnam
Yemen, Rep.
Zambia
Zimbabwe
Total fertility rate
Year
1940
1960
1980
2000
2020
2
3
4
5
6
7
8
9
5.5
6.5
7.5

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