Growth and Childbearing in the Short- and Long-Run*

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Abstract

We analyze economic growth and fertility change in the developing world over six decades, using data on 2.3 million women from 255 surveys in 81 countries. We find that fertility responds differently to fluctuations and long-run growth, and the nature of these responses varies over the lifecycle. Fertility is procyclical, falling during recessions, but also declines with long-run growth. Lifetime fertility is affected by fluctuations near the end of the reproductive period, but not those at prime reproductive age. Our results are consistent with models linking demography, human capital, and long-run growth, extended to include a lifecycle with liquidity constraints.

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1 Introduction

The relationship between macroeconomic conditions and fertility is sometimes positive and sometimes negative. In historic pre-industrial populations, birth rates rose during eras of relative prosperity (Galloway 1988; Lee 1997), much in line with Malthus's characterization of fertility control as a "preventive check" on population size (Malthus 1798). Birth rates are similarly procyclical in present-day industrialized countries, as evidenced by a voluminous literature reviewed by Sobotka, Skirbekk, and Philipov (2011). But fertility is currently highest in the poorest and least-developed countries, suggesting that the economies that grew most in recent history contain the populations in which reproduction slowed most (Bongaarts and Watkins 1996).[1] The mixed empirical record impedes theoretical progress on the linkages between economic growth and the demographic transition.

This paper empirically characterizes the relationship between economic growth and fertility change during the process of economic development, with an eye toward reconciling the wide range of aggregate elasticities one can find in the data. The key insight is that fertility may respond differently to growth over different time horizons. Long-run growth may alter prices and returns in ways that short-run fluctuations do not, while short-run fluctuations may have liquidity and intertemporal substitution effects that are absent in the long run. A rough way to conceptualize this difference is to think of long-run growth as influencing the ideal number of children and of fluctuations as influencing the timing of childbearing. For example, unified growth theory (Galor 2011) emphasizes how long-run growth may reduce the optimal number of children by increasing the return to child investment or women's work, forces unlikely to play out over the business cycle or even perhaps within an individual lifecycle.[2] Similarly, models of fertility over the lifecycle (Hotz, Klerman, and Willis 1997) highlight how income or price fluctuations can alter the timing of childbearing, even if lifetime fertility remains constant. To take both perspectives into account, we consider short-run and lifecycle fertility responses to fluctuations alongside longer-run relationships.

For studying growth-fertility linkages at different time horizons, we combine macroeconomic data with fertility data on 2.3 million women from 255 World Fertility Surveys (WFS) and Demographic and Health Surveys (DHS), covering 81 low- and middle-income countries from 1950 to 2010. The

[1] Further complicating the record, Myrskylä, Kohler, and Billari (2009) confirm that fertility is currently highest in the least-developed countries but also find a positive development-fertility relationship in the most-developed countries.

[2] The idea that technological progress raises the return to human capital dates back to Nelson and Phelps (1966). In an separate but related line of work, Barro and Becker (1989) show that at a given interest rate, higher productivity growth lowers fertility by incentivizing higher consumption growth.

survey data allow us to avoid standard cross-national databases on fertility, which rely heavily on interpolation, smoothing, and demographic modeling (United Nations 2015; World Bank 2015)—especially problematic for the study of fluctuations. We also use the survey data to calculate cohort measures of fertility and investigate within-country heterogeneity—difficult to do with cross-national databases—and to better match the timing of conception (rather than birth) to fluctuations.

With these data in hand, we can provide a geographically and temporally broader account of the growth-fertility relationship at various time horizons than has previously been possible. In fact, estimates of this relationship in contemporary developing countries are surprisingly rare at *any* time horizon. A small literature has documented drops in fertility during economic crises in select developing countries, consistent with the procyclicality observed in other settings, but this research does not speak to the distinction between the short and long run, and its generalizability to a broader set of countries is not clear.[3] Analyses of growth and fertility change across countries over longer time horizons, though less plagued by the limitations posed by interpolated fertility indicators, are less conclusive and even rarer.[4] Also unknown in developing countries is whether any fertility responses to short-run fluctuations extend beyond adjustments in a woman's childbearing timing to affect her lifetime number of children, as they do in the United States (Currie and Schwandt 2014).

We address these gaps in the literature with three analyses: two on the flow of fertility at the population level (i.e., period fertility) and one on the stock of fertility at the cohort level (i.e., cohort fertility). First, we examine how the annual rate of starting a successful pregnancy (i.e. conceiving a future liveborn child) responds to growth fluctuations in the short run, paying attention to whether immediate effects are subsequently offset and whether fertility responds differently to booms and busts. Second, we estimate how the long-run rate of economic growth relates to the long-run rate of fertility decline. Third, we look across birth cohorts within a country to ask how completed fertility varies with macroeconomic conditions experienced at different phases of the reproductive lifecycle.

To motivate these empirical analyses, we explore a lifecycle extension of the parental choice problem from the theoretical literature on long-run growth and the demographic transition, in which liquidity-constrained parents with uncertain incomes and wages make consumption, childbearing, and

[3]See National Research Council (1993); Tapinos, Mason, and Bravo (1997); Lindstrom and Berhanu (1999); and Adsera and Menendez (2011), which with 18 Latin American countries provides the broadest geographic coverage.

[4]Bongaarts and Watkins (1996) find that Human Development Index changes negatively predict total fertility rate changes from 1960 to 1990. Schultz (1997) finds that, conditional on covariates, economic growth was uncorrelated with total fertility rate changes from 1972 to 1988. Murtin (2013) finds a non-linear relationship for 1870-2000 but does not distinguish the short from long run.

child investment choices over a finite reproductive period.[5] This theoretical framework highlights how procyclical fertility can coexist with a negative long-run relationship between economic growth and fertility change, as well as how income or wage fluctuations toward the end of the reproductive period can have permanent effects on fertility, while the effects of earlier fluctuations may be offset by subsequent adjustments in childbearing decisions.[6] In the framework, optimal fertility *timing* is heavily influenced by liquidity constraints and intertemporal labor substitution over the lifecycle, while optimal *lifetime* fertility is heavily influenced by long-run growth and its implications for wage rates or the quality-quantity tradeoff.

Our empirical results illustrate these implications of the lifecycle framework. When we study annual fluctuations, we find that fertility is procyclical within the reproductive lifecycle, with statistical significance at prime childbearing ages. The procyclicality is driven entirely by downturns and is stronger for less-educated women, reinforcing the potential role of liquidity constraints that prevent poor households from smoothing through recessions. In contrast, when we study average rates of change over periods of 20 or more years, we find that fertility declines with long-run economic growth. More specifically, prime-age fertility declines more rapidly in faster-growing countries, while older-age fertility declines more slowly; on net, the former effect dominates that latter for total fertility. These patterns reflect a simultaneous decline and delay of fertility in faster-growing countries, which can also be reconciled with a lifecycle model in which long-run growth raises the return to human capital investment and women's work. When we compare cohorts within a country, we find that growth experienced in the prime childbearing years is unrelated to lifetime fertility, while growth in the 30s leads to higher lifetime fertility—again consistent with our theoretical predictions regarding offset opportunities early and late in the reproductive lifecycle. Growth fluctuations thus affect both the tempo (timing) and quantum (lifetime cumulation) of fertility (Bongaarts and Feeney 1998). We show that many of our findings are obscured in standard cross-national fertility data, while many others are altogether impossible to study, underscoring the benefits of combining hundreds of surveys.

The short- and long-run analyses of period fertility differ substantially in their statistical interpretation, in that the short-run analysis uses a demanding regression specification that isolates within-country, within-age, and within-year variation, while the long-run analysis must rely on cross-

[5] See Galor and Weil (2000); Galor and Moav (2002); Hazan and Berdugo (2002); De La Croix and Doepke (2003); and Cervellati and Sunde (2015).

[6] See Foster (1993) for a related model of how short- and long-run economic-demographic relationships can differ.

country (although still within-age) variation in rates of change. Indeed, the short-run results are robust to including a range of fixed effects, trends, and covariates, suggesting that they reflect a causal effect of aggregate fluctuations. Conversely, causality is harder to infer from the long-run results, which reflect correlations between long-run growth and other trends that incentivize lower fertility. Notably, these results are not explained by the initial levels of GDP per adult and population density; nor by rates of change in adult female education, adult female labor force participation, the sectoral composition of value added, urbanization, infant mortality, conflict, and democratization.[7] Instead, we find only one variable that explains a meaningful part of the negative association between long-run economic growth and fertility change: the rate of change in secondary school enrollment (among teens, not mothers). Because rising enrollment suggests rising returns to human capital investment, these results are consistent with human capital-based theories of unified growth.[8]

While our findings shed light on the evolution of fertility during the process of growth (and recession), they do not directly speak to whether the demand for children exhibits positive wealth or income effects. That question, which has long interested economists, is related to our investigation, but substitution and liquidity effects are key to understanding our short- and long-run results. In this sense, our contribution should be seen as separate from studies identifying positive wealth or income effects on fertility using variation from natural resource booms within or across countries (Lovenheim and Mumford 2013; Brueckner and Schwandt 2015) or from housing booms in the United States (Lovenheim and Mumford 2013; Dettling and Kearney 2014).[9] Theoretically and methodologically, our analysis is more similar to research on the link between economic growth and mortality decline at various time horizons (Deaton 2007; Baird, Friedman, and Schady 2011), as well as on the cumulative mortality effects of economic shocks over the lifecycle (Cutler, Huang, and Lleras-Muney 2016).

We focus on how economic growth and its correlates affect fertility, rather than causality running the other way, from fertility to growth. The idea that high fertility causes underdevelopment has long concerned researchers and policymakers (Coale and Hoover 1958); recent findings support this concern but suggest that the effects are modest in size (Ashraf, Weil, and Wilde 2013; Miller and

[7]The ideal test for the role of women's work would use wages (Schultz 1985; Heckman and Walker 1990), not labor force participation or sectoral composition. Unfortunately, wage data are unavailable for most of our sample.

[8]A more directly relevant covariate would be parents' expected return to human capital investment at the time of the childbearing decision, but this variable is unavailable. Contemporaneous school enrollment is an imperfect proxy.

[9]In a related study that spans the literatures on both income effects and the business cycle, Schaller (2016) interacts nationwide industry-level labor demand shocks with the local industrial composition of employment to identify the fertility effects of shocks to men's and women's labor market opportunities in the United States.

Babiarz 2016). Though this concern is also relevant for models of economic growth and the demographic transition, it is beyond the scope of this paper, so we rule out any first-order effects in this direction by studying growth in GDP per adult rather than GDP per capita. Most specifically, our results pertain to the relationship between economic growth and fertility change after an economy escapes the Malthusian trap but before it attains high income per capita and low fertility. In the Malthusian era, although macroeconomic fluctuations had similar effects on fertility (Galloway 1988; Lee 1997), sustained productivity growth led to higher population density rather than rising living standards and declining fertility (Ashraf and Galor 2011). In the modern era, the balance of income and substitution effects from long-run growth may differ. Indeed, when we examine data from contemporary industrialized countries as a complementary exercise, we find similar procyclicality but no association between growth and fertility change in the long run. Thus, our paper clarifies the forces underlying growth-fertility linkages at various horizons, both over time and over the lifecycle, for researchers and policymakers interested in demographic aspects of the development process.

2 Lifecycle Considerations in Models of Demography and Growth

In the standard overlapping generations formulation, models of long-run economic growth and the demographic transition posit parents making choices regarding the quantity (and sometimes quality) of their children in a single period. Though tractable, this setup precludes a distinction between short- and long-run forces, as well as any consideration of the timing of fertility over the lifecycle.[10] In this section, we build on these models to explore how fertility behavior responds to fluctuations and long-run growth over a finite reproductive lifecycle. Our key goal is to understand how the balance of income and substitution effects, as well as their persistence, varies across different time horizons and different points in the lifecycle.

As is common in this literature (Jones, Schoonbroodt, and Tertilt 2010), we model parents' utility as separable over their own consumption c_t, their number of children n_t, and their children's mean human capital h_t:

$$U\left(c_t, n_t, h_t\right) = u_c\left(c_t\right) + u_n\left(n_t\right) + u_h\left(h_t\right) \tag{1}$$

[10] Doepke, Hazan, and Maoz (2015) and Jones and Schoonbroodt (2016) calibrate overlapping generations models with lifecycles and endogenous fertility but assume either deterministic income or a single childbearing period. A microeconomic literature reviewed by Hotz, Klerman, and Willis (1997) estimates lifecycle fertility models that rarely include long-run forces and make restrictive credit market assumptions. Sommer (2016) calibrates a lifecycle fertility model with liquidity constraints but studies the effects of earnings risk, rather than fluctuations or long-run growth.

However, we treat t as a period rather than a generation, so that $U(\cdot)$ becomes a period utility function. Specifically, we consider a generation of parents who live $t = 1, \cdots, T$ periods and can bear children until they reach menopause at age $M < T$. We assume that each sub-utility function $u_x(\cdot)$ is increasing, concave, and twice continuously differentiable, with $\lim_{x \downarrow 0} u'_x(x) = \infty$ and $\lim_{x \uparrow \infty} u'_x(x) = 0$. Parental choices maximize expected utility, discounted by factor β.

Parents start their lives with assets A_0 and then receive stochastic wages w_t (with a period time endowment of 1) and unearned income y_t in subsequent periods. Because w_t factors into the time cost of children, but y_t does not, one can also think of these variables as reflecting women's and men's labor income, respectively. Some models include subsistence consumption constraints to accommodate both Malthusian and modern population dynamics, but we do not, since we are mainly interested in fertility behavior after the emergence of sustained economic growth. However, we do impose a borrowing constraint ($A_t \geq 0$), which may be key to understanding the effects of fluctuations.[11] In each period, parents allocate their potential income $w_t + y_t$ among their own consumption, the quantity and quality costs of children, and savings (which earn gross return R).

Although the most realistic setup would treat births as binary, we allow parents to set a birth rate $b_t \in [0, 1]$ in each period.[12] Parents begin their lives with no children and accumulate them according to: $n_t = n_{t-1} + b_t$. To avoid keeping track of children's ages, we assume that parents pay child costs only once, at birth. Each child costs $\tau \in (0, 1)$ units of time and κ units of the consumption good, plus any education spending e_t to produce human capital.[13] Education spending is transformed into human capital by a (twice continuously differentiable) human capital production function $h(e_t; \bar{g})$, which we allow to depend on the long-run growth rate of technology $\bar{g}$. Following Galor and Weil (2000), we assume that education increases human capital at a decreasing rate ($h_e > 0$, $h_{ee} < 0$) while long-run growth depletes human capital but makes education more productive ($h_{\bar{g}} < 0$, $h_{e\bar{g}} > 0$).

Because our primary interest is not the allocation of education within the family, we simplify by assuming that parents plan a single education level e for all of their children in period 0 of the model, before the first period of the lifecycle. From period 1 to period $M - 1$, parents then make a sequence of consumption and birth decisions, followed by a sequence of consumption decisions from menopause

[11] By modeling an endowment economy with aggregate shocks in general equilibrium, Jones and Schoonbroodt (2016) can generate procyclicality (over periods much longer than those considered here) without borrowing constraints.

[12] Besides following much of the literature, continuous fertility allows us to derive first-order conditions that capture much intuition while avoiding the complications of discrete/continuous dynamic models (Iskhakov et al. 2015).

[13] The assumption that education has no time cost serves only to avoid corner solutions in which education exhausts the parents' period time constraint. It does not substantively affect the analysis.

M to death T. In light of the finite horizon, we work backward: characterizing the consumption sequence first, the birth sequence next, and education spending last.

The first-order conditions to the lifecycle problem lead to a standard consumption Euler equation:

$$u_c'(c_t) = \beta R E_t \left[u_c'(c_{t+1})\right] + \lambda_t \tag{2}$$

where λ_t is the Lagrange multiplier on the borrowing constraint. When the borrowing constraint does not bind, parents set the current marginal utility of consumption to the discounted expected marginal utility of consumption in the next period. When it does bind, they fall short of consuming enough in the current period to satisfy this condition, with a positive multiplier filling the gap. Although the Euler equation does not directly involve fertility, the consumption smoothing motive is key to understanding the timing of births over the business cycle. This point becomes apparent upon inspection of a separate first-order condition, which equates the marginal benefit of consumption with the discounted marginal benefits of childbearing:

$$u_c'(c_t) \quad = \quad \frac{u_n'(n_{t-1}+b_t) + \sum_{s=t+1}^{T} \beta^{s-t} E_t\left[u_n'(n_s)\frac{\partial n_s}{\partial b_t}\right] + \mu_t^0 - \mu_t^1}{\tau w_t + \kappa + e} \tag{3}$$

where μ_t^0 and μ_t^1 are the Lagrange multipliers on the constraints that the birth rate be at least 0 and at most 1, respectively.

Equations (2)-(3) provide much insight into the effects of wage and income fluctuations. Because the right-hand side of equation (3) is divided by the marginal cost of childbearing—which includes w_t—a temporary fall in the wage incentivizes the shifting of births from the future to the present. If T is large relative to M, then fluctuations in w_t before menopause have minimal effects on expected lifetime income, so this intertemporal substitution effect is likely to dominate any income effects when the borrowing constraint does not bind. Under these conditions, fluctuations in y_t are also unlikely to have a large effect on the birth rate when T is large. When the constraint binds, however, a negative wage or income shock decreases current consumption by equation (2), which then also incentivizes reduced childbearing by equation (3).[14] The possibility that childbearing behavior is procyclical due to borrowing constraints also implies possible non-linearity and heterogeneity. Borrowing constraints

[14] Existing research on fertility over the business cycle has also noted offsetting liquidity and intertemporal substitution effects, albeit generally without fully specifying a model (e.g., Ward and Butz 1980; Adsera and Menendez 2011).

may bind more during deep recessions, and the poor may be especially likely to hit zero assets.[15]

The durability of children adds several nuances in equation (3).[16] The lagged number of children appears inside $u_n'(\cdot)$, placing a ceiling on the marginal utility of children for all $t > 1$. No such ceiling exists for the marginal utility of consumption, implying that births decline to zero more rapidly than consumption when parents are borrowing constrained. Also because the lagged number of children appears inside $u_n'(\cdot)$, parents may offset past adjustments in childbearing. For example, if a negative shock forced borrowing-constrained parents to forego births in period $t-1$, then the marginal gains from childbearing are high in period t. Demographers refer to intertemporal substitution of births of this type as a tempo effect.[17] When offset is incomplete ($\frac{\partial n_s}{\partial b_t} > 0$ for $s > t$), the benefit of current childbearing includes the marginal utility of children in the future. In fact, offset becomes impossible after menopause, so wage and income shocks toward the end of the reproductive lifecycle may be more likely to have permanent effects on the number of children. As a result, parents approaching menopause may tolerate greater declines in consumption to finance childbearing.

The insights about fluctuations from equations (2)-(3) do not necessarily carry to long-run productivity growth, which may be built into parents' expectations and may alter prices and returns in ways that fluctuations do not. In general, we are agnostic about how long-run growth affects the parents' budget constraint, but as a starting point, it is useful to assume that $\bar{g}$ reflects the growth rate across generations but does not affect wages or incomes in the parents' lifetimes. This approach mimics overlapping generations long-run growth models and allows us to focus on substitution effects from $\bar{g}$. Assuming an interior solution, the education spending plan satisfies the first order condition:

$$u_h'\left(h\left(e;g\right)\right)h_e\left(e;g\right)\left(\tfrac{1-\beta^T}{1-\beta}\right) = \sum_{t=1}^{M-1}\beta^{t-1}E_0\left[\nu_t b_t\right] \tag{4}$$

where ν_t is the multiplier on the budget constraint in period t. Higher $\bar{g}$ raises the left-hand side, which leads parents to rasie e. The increase in e pushes up the denominator in equation (3), so the optimal number of children declines (which then reinforces the increase in e in equation [4]).

Thus, a central mechanism of unified growth theory—in which technology growth causes lower fertility by increasing the return to child investment—remains intact, even in a model that admits

[15]This last claim comes up against the reality that fertility control is imperfect, which we have not included in the model. If the poor are also less effective at regulating the timing of births, then the prediction becomes ambiguous.

[16]These insights build on Becker (1960), who first introduced the concept of children as durable goods.

[17]In the very short run, some of the tempo effect will be biological due to postpartum infecundity.

procyclical fertility in the short run. This prediction relates long-run economic growth to the the *level* of fertility, while many of our analyses examine *changes* in fertility, but this gap is easy to address with slight adjustments to the theory. For example, if parents adapt to the new economic environment slowly, then long-run growth will be associated with long-run (rather than instantaneous) fertility decline. A similar prediction obtains if higher $\bar{g}$ incrementally raises the return to education spending from its most recent value.

If $\bar{g}$ affects parents' budget constraint, then these predictions become less sharp. In this case, the effect of long-run growth on fertility additionally depends on the balance of income effects from higher w_t and y_t and substitution effects from higher w_t. Many economic theories of the demographic transition emphasize the link between long-run growth and rising women's wages (Schultz 1985; Galor and Weil 1996), which in our framework can be seen as a shift from y_t to w_t in the composition of household income. By equation (3), such a shift incentivizes fertility reduction. If this mechanism is at work, then long-run growth will also cause fertility to become more countercyclical.

For parsimony, our framework abstracts from several additional margins that may be relevant to fertility change. First, we ignore human capital investment in adulthood, a margin that may imply additional lifecycle fertility consequences of long-run growth. Suppose parents could supplement their own human capital by investing time in further schooling or on-the-job training, and suppose long-run growth also raised the returns to these activities. A standard result is that individuals front-load human capital investment at young ages to reap returns over a longer horizon (Ben-Porath 1967). By this reasoning, long-run growth leads to fertility *delay* in addition to *decline*.[18] Second, we do not directly consider the proximate determinants of fertility that underlie fertility responses to growth. Intentional changes in childbearing, as captured in the framework, may be achieved through some mix of traditional birth control (including marriage timing), modern contraception, and abortion.[19] Other proximate determinants that may act as mediators, such as unintentional changes in coital frequency (e.g., from stress or migration) or infecundity (e.g., from malnutrition), are less natural in our choice framework but still relate to its emphasis on liquidity constraints and intertemporal substitution. Our empirical analysis examines marriage timing but not these other mechanisms because of data limitations. Third, we do not explicitly model child mortality. In our empirical analysis, we find no evidence that child mortality explains our results.

[18]See Blackburn, Bloom, and Neumark (1993) for a model making this point.

[19]Alam and Pörtner (2016) find that crop losses increase modern contraceptive use and decrease fertility in Tanzania.

3 Data

Our analysis requires careful measurement of economic growth and fertility. For the former, we obtain data on GDP from the Penn World Table (PWT) v. 8.1 (Feenstra, Inklaar, and Timmer 2015), matching it with data on population and age structure from the United Nations. The central independent variable is the logarithm of GDP per adult age 15-64 for country c in year t, $GDPpa_{ct}$, where the age range for the denominator is chosen to minimize concern about the endogeneity of population size due to endogenous fertility and mortality. To ease interpretation, this variable is multiplied by 100, so that the results are quantified in log points, which approximately reflect percentage points. When analyzing levels, we adjust for purchasing power parity (PPP); when analyzing growth rates, we adjust for inflation but use national prices, following the recommendation in Johnson et al. (2013). Our focus on growth in production rather than technology follows the empirical literature but departs somewhat from the theory.

To measure fertility, we assemble data from all publicly available, standard format WFS and DHS surveys that are nationally representative for all women and can be merged with our macroeconomic dataset, leading to a sample of 81 countries. Appendix Table A1 lists the number of surveys for each of these countries, which were all classified as low- or middle-income at the time of the surveys (World Bank 2015). In both the WFS and the DHS, respondents provide full birth histories, listing all of their children ever born, with information on birth date and survival status.[20] These data allow us to track fertility behavior over time and over the lifecycle, although they are sometimes subject to reporting errors (Schoumaker 2014), a matter we discuss further below.[21] Some surveys only interviewed women who had ever been married or had completed schooling; for these surveys, we only use data on women who at the time of the survey belonged to an age group in which the rate of ever marriage or school completion exceeded 95%.

For all analyses, we collapse the individual-level data into country-year-age or country-cohort cells, allowing us to weight countries in a consistent way across a range econometric models. To estimate the fertility rate for each cell, we pool data from all surveys in the same country and rescale the survey weights to reflect each survey's sample size contribution to the cell.[22] We exclude cells

[20] The DHS from El Salvador 1985 and Nigeria 1999 have well-known deficiencies in their birth histories (Casterline and Odden 2016). For these surveys, we do not use the birth histories but do use data on lifetime fertility.

[21] Reporting errors take the form of both omitted births (which are likely to have been in the distant past or to have involved deceased children) and displaced births (either forward or backward in time).

[22] Motivated by concerns that reporting errors are more likely for longer recall periods, we apply a Bartlett kernel to

with fewer than 30 observations (<5% of cells) for precision. We generate two types of fertility rates: period, summarizing fertility outcomes across ages in a given year, and cohort, summarizing the lifetime fertility outcomes of women born in the same year.

For the analysis of period fertility, we study the age-specific conception rate, CR_{cta}: the number of conceptions per 1000 women aged a in year t from country c. Because we do not have information on miscarriages or abortions, we focus only on conceptions that resulted in a live birth; because we do not have information on gestational age at delivery, we assume that conception took place 9 months before the date of birth.[23] As such, we do not directly analyze childbearing behavior but instead use approximate dates of conception for fetuses that survived gestation: a limitation, given that economic conditions may affect the probability of fetal death. The term "conception rate" is thus a slight abuse of demographic terminology, but one that follows Currie and Schwandt (2014) in their leading work on economic conditions and fertility. This choice to focus on the timing of conception rather than birth also follows Currie and Schwandt, under the reasoning that economic growth may be more likely to affect the former.

To help distinguish between the short and long run, we focus on country-age combinations with conception rates and macroeconomic data spanning at least 20 years. As reported in the first two columns of Table 1, this sample definition gives rise to 58,988 distinct cells defined by country, year, and age, with conception rates based on the fertility histories on 2.3 million women from 65 countries. Pooling all ages 15-44, age-specific conception rates have a mean of 199 per 1000. The annual change in log GDP per adult, which approximates the annual growth rate, averages 1.0 log points, with a standard deviation of 5.8. Female education averages 4 years; the urbanization rate, 38 percent; and GDP per adult, 4,476 international dollars, adjusted for purchasing power parity.

For the analysis of cohort fertility, we study the completed fertility rate, CFR_{cj}: the number of children per 1000 women from country c and birth cohort j. We only include women over 45 at the time of the survey, treating their fertility as complete. Our main cohort analyses are based on all children ever born, although we show that we obtain similar results when we only count children who survived to the date of the survey.

As reported in the last two columns of Table 1, data are available on 935 country-cohort cells from

the rescaled survey weights in a robustness exercise, down-weighting births that occurred long before the survey.

[23]We count multiple births as coming from a single conception and allow for the possibility that a woman may conceive twice in one year.

62 countries, containing 212k women over 45. The completed fertility rate per 1000 women averages 5951 children ever born and 4862 surviving children; an individual woman experiencing these rates would bear 6 children, of whom 1 would die before she reached her late 40s. Compared with the period sample, the cohort sample is characterized by a higher average age because it excludes women below age 45. Average educational attainment is also lower because earlier cohorts received less education. Other characteristics have similar means in the two samples.

For covariates, analyses of heterogeneity, alternative measures of fertility, and comparisons with developed countries, we draw on several additional aggregate data sources. We obtain alternative fertility data from the World Development Indicators (WDI) and United Nations (UN); information on contraceptive use, population density, and the sectoral composition of value added from the UN; school enrollment and labor force statistics from the WDI and International Labor Organization (ILO); democratization scores from the POLITY IV project; and conflict indicators from the UCDP/PRIO Armed Conflict Dataset. Other covariates, like average female education, the infant mortality rate, and the urbanization rate, are estimated from the WFS/DHS microdata. For comparison with developed countries, we use the Human Fertility Database, which brings together vital registration data from wherever they are of high quality.

4 Analysis of Period Fertility

Our analysis of period fertility focuses on how changes in fertility across the age distribution vary with short- and long-run economic growth. We begin by laying out how we distinguish between short- and long-run patterns empirically, followed by the main results for both horizons. We then delve further into the results—assessing non-linearity, lag structure, heterogeneity, and alternative covariates—first for the short run and then for the long run.

4.1 Defining Time Horizons

A key issue is how to define "short run" and "long run." To allow the data to speak to this issue, we run a series of first-difference regressions in which we vary the length of the difference. For each 5-year age group A from $[15, 19)$ to $[40, 44)$, we run:

$$CR_{cta} - CR_{c,t-\Delta,a} = \beta^{A} \left(Y_{ct} - Y_{c,t-\Delta}\right) + \alpha_a^{A} + \varepsilon_{cta}^{A} \quad (5)$$

where $Y_{ct} = 100 \times \ln(GDPpa_{ct})$. Because the distribution of ages within each age group varies across countries and over time, we include a single-year age effect α_a^A, thus allowing fertility levels to trend differently for each age within the age group. We estimate equation (5) using a range of values for the length of the difference Δ, from 1 to 30 years. Figure 1 displays the results, for each age group plotting estimates of the coefficients against the length of the difference.

The patterns in Figure 1 make clear that economic growth and fertility change have different relationships over different time horizons. In the annual first difference ($\Delta = 1$), all age groups have positive coefficients, indicating procyclical fertility. However, all but one age group immediately begin trending downward with rising Δ, becoming negative by $\Delta = 14$ and leveling off at about $\Delta = 20$. In other words, for all age groups less than 40, economic growth over periods of 14 or more years is negatively associated with fertility change over the same period. Meanwhile, above 40, economic growth is positively related to fertility change at all time horizons. These patterns have several noteworthy implications, both statistical and economic. First, the associations at longer time horizons bias the annual first-difference coefficients estimated with equation (5), with the magnitude of the bias depending on the relative contributions of transitory fluctuations and long-run growth to the variance of the annual growth rate. For all ages under 40, the long-run associations are negative, implying that a naive first-difference specification understates the procyclicality of conceptions; above 40, the positive long-run association implies the opposite. Any analysis of short-run procyclicality must correct for these biases. Second, the sharp drop of the coefficients for the first five age groups beyond $\Delta = 1$, as well as their leveling at about $\Delta = 20$, suggest 1 and 20 years as reasonable definitions of the short and long run. Third, the implied procyclicality with varied long-run associations is consistent with the theories described in Section 2, although the subsequent analysis will need to address robustness to covariates and statistical significance.

4.2 Methods

Based on the patterns in Figure 1, we divide our study of period fertility into a short-run analysis of annual fluctuations and a long-run analysis of average annual changes over periods of at least 20 years. To distinguish time horizons as clearly as possible, we restrict both analyses to country-age cells that span at least 20 years. For the short-run analysis, we modify equation (5) to eliminate bias from the long-run relationship by including a country effect, which absorbs the country's long-run average rate of change in GDP per adult and the conception rate. For completeness, we also include a year effect, addressing any concern about spurious global trends. The first-difference specification then becomes:

$$\Delta CR_{cta} = \beta^A g_{ct} + \lambda_c^A + \tau_t^A + \alpha_a^A + \varepsilon_{cta}^A \tag{6}$$

where ΔCR_{cta} is the change in the conception rate from the previous year, and g_{ct} is the annual change in $100 \times \ln(GDPpa_{ct})$, which approximates the growth rate. λ_c^A, τ_t^A, and α_a^A are the country, year, and age effects, which in first differences serve to control for correlated level trends.[24] The coefficient β^A isolates how fluctuations of the growth rate from its long-run country average affect changes in the conception rate, net of year- and age- specific factors. A 1-log point growth fluctuation raises the change in the conception rate by β^A.

For the long-run analysis, we seek to estimate a long-difference version of equation (5). The standard approach would relate simple changes in log GDP per adult to simple changes in fertility over an interval of 20 years. However, because our country-year-age conception rates are noisy estimates from surveys, we deviate from this standard approach in order to leverage as much information as possible on the rate of long-run fertility change. Instead of the long difference, we analyze the average annual rates of change in the two variables, $\overline{g}_{ca}$ and $\overline{\Delta CR}_{ca}$, over periods of at least 20 years. To estimate these quantities using as much information as possible, we regress $100 \times \Delta \ln(GDPpa_{ct})$ and CR_{cta} on year within each country-age cell, using the slope of the trend as the estimated average annual rate of change.[25] We then run the following regression for each 5-year age group A:

$$\overline{\Delta CR}_{ca} = \beta^A \overline{g}_{ca} + \alpha_a^A + \varepsilon_{ca}^A \tag{7}$$

As before, the single-year age effect α_a^A absorbs any age-related factors common across countries. By collapsing the country-year-age observations into country-year cells, we remove the time dimension from our panel, so the year effect τ_t^A drops out. Similarly, because equation (7) primarily analyzes

[24] Equation (6) can be obtained from differencing a level specification with country (μ_c^A), year ($\tilde{\tau}_t^A$), and age (ω_a^A) fixed effects, as well as country (λ_c^A) and age (α_a^A) linear trends:

$$CR_{cta} = \beta^A \left(100 \times \ln(GDPpa_{ct})\right) + \mu_c^A + \tilde{\tau}_t^A + \omega_a^A + \lambda_c^A t + \alpha_a^A t + \tilde{\varepsilon}_{cta}^A$$

On differencing, μ_c^A and ω_a^A drop out, while λ_c^A, $\tau_t^A \equiv \Delta\tilde{\tau}_t^A$, and α_a^A become country, year, and age effects. However, serial correlation, non-stationarity, and the need for PPP adjustment make the level specification unattractive.

[25] An alternative approach would simply take the mean of observed annual changes g_{ct} and ΔCR_{cta}, but this approach would not use all available information due to gaps in the data. For example, if data were collected only in 1970, 1971, 1990, and 1991, then the mean of the two observed annual changes would ignore developments during 1971-1990.

variation in ΔCR_{cta} and g_{ct} that was absorbed by the country effect in equation (6), we omit λ_c^A from the long-run regression. Here, β^A represents the cross-country association of long-run economic growth with long-run fertility change, net of age-related factors. A country with a 1-log point faster rate of long-run growth fluctuation experiences a β^A.

For both equations (6) and (7), we summarize the age group results by reporting the implied result for the total conception rate (TCR) per 1000 women, defined as the expected number of conceptions in a hypothetical cohort of 1000 women who experience current age-specific conception rates at every age from 15 to 44:

$$\beta^{TCR} = 5\left(\sum_A \beta^A\right) \tag{8}$$

This summary measure is the sum of the age group coefficients, multiplied by 5 to account for the length of each age group. While heterogeneity over the age distribution is key to our investigation, β^{TCR} provides an overall measure of the association between economic growth and fertility change.

To assess the roles of other aggregate variables in explaining any relationship between economic growth and fertility change, we report estimations of equations (6) and (7) with and without a main set of covariates. In the extended models, we control for the initial level of population density and log PPP-adjusted GDP per adult, as well as the change (for the short-run analysis) or average rate of change (for the long-run analysis) in female education, urbanization, infant mortality, and armed conflict.[26] These variables serve as our main controls because they may be key drivers of demographic change, and they are available for all country-years in our dataset. We consider covariates that are available only for subsamples later.

Female education and urbanization are averaged at the country-year-age level, just as with the conception rate. Because these characteristics are measured at the time of the survey, they change minimally over the lifecycle for women born in the same country and year. Thus, when we compute changes from years $t-1$ to t in the average characteristics of women aged a, we effectively measure differences between birth cohorts $c, t-a-1$ and $c, t-a$.[27] Fluctuations in this cohort difference are unlikely to be correlated with fluctuations in growth over time, so these variables pose little omitted variable risk in the short-run analysis; nevertheless, we include them for completeness.[28]

[26] Some of our these covariates may mediate an effect of growth, leading to risk of over-controlling.

[27] Some changes may occur over the lifecycle if women born in the same country and year are surveyed multiple times, but such changes are generally small in the data.

[28] At younger ages, education and sector of residence may respond to growth fluctuations, raising concern that reverse causality in the covariates may bias estimates of β^A.

Analogously, when we compute average annual rates of change for a specific age over multiple years, we effectively measure trends across birth cohorts. Because these trends may be correlated with long-run growth, female education and urbanization may be more relevant in the long-run analysis.

The remaining covariates may be important in both analyses. For infant mortality, we pool all maternal ages to measure rates at the country-year level in an effort to minimize noise. In the short-run analysis, we track deaths among infants conceived in the previous rather than current year, both because these deaths are less likely to be endogenous and because they are more likely to influence conceptions in the current year.[29] Thus, we control for the change in the overall infant mortality rate from year $t-2$ to $t-1$.

For weighting and variance estimation, we make conservative choices that clarify interpretation. Sample sizes in individual WFS and DHS surveys range from fewer than 5,000 to more than 100,000 women, suggesting possible efficiency gains from weighting cells by their sizes. We choose to weight cells equally to ease interpretation of the results, effectively giving countries balanced representation within each age group or birth cohort. When relevant, we show in the Appendix that weighted analyses lead to similar results. We also cluster standard errors by country, allowing for arbitrary error covariance within country while imposing independence across countries.

4.3 Main Results

Table 2 reports the main results of the analysis of period fertility. The first two columns summarize the level of fertility and its rate of change. In column (1), average conception rates follow an inverted u-shape in age, peaking at 261 per 1000 among 20-24 year olds. The TCR per 1000—the number of conceptions expected of a hypothetical cohort of 1000 women experiencing these age-specific conception rates over their reproductive lifecycles, is 5483, or 5.5 conceptions per woman. Despite this high level, fertility was falling throughout the sample period, as revealed in column (2). On average, conception rates in all age groups declined by 1 to 3 points per year, with a -67 annual change in the TCR, corresponding to a decadal reduction of two-thirds of a conception per woman.

These rates of fertility change serve as dependent variables in the remainder of Table 2. Columns (3)-(4) report the short-run coefficients from equation (6), first without the main covariates and

[29] We include an indicator for missing mortality information in the extended short-run model, mainly to deal with the absence of lagged mortality data for the first cell in any country-age series. Results are unchanged if we drop these cells instead.

then with them. In the basic model without covariates, all of the short-run coefficients are positive, indicating procyclical fertility, with statistical significance in the prime childbearing ages. The largest coefficient of 0.56 implies that a one log point increase in GDP per adult raises the number of conceptions by roughly $\frac{1}{2}$ per 1000 25-29 year olds. Moving from 25-29 to neighboring age groups, the coefficients decline more than would be proportional to the level of fertility. This finding is consistent with the theoretical framework's prediction that older parents (who are closer to menopause) are less willing to forego births during a recession, although it is only suggestive evidence. Combining all age groups, the TCR increases by 8.8 per 1000 in response to a log point positive growth fluctuation. As reported in Table 1, the annual change in log GDP per adult has a standard deviation of 5.8, which puts these magnitudes in further perspective. The standard deviation shrinks only slightly, to 5.4, after conditioning on the fixed effects in equation (6), so one can multiply the results by 5-6 to gauge the effect of a one standard deviation growth fluctuation.

Alternative regression specifications leave these results intact. In column (4), the addition of covariates in the extended model does not meaningfully alter the coefficients or their significance levels. For completeness, we report the coefficients on these covariates in Appendix Table A3. Changes in female education, urbanization, and infant mortality all have significant relationships with conception rate changes at various ages, yet none meaningfully explains the procyclicality of conceptions. Additional robustness checks are reported in Appendix Figure A1, which plots age-group-specific coefficients from a range of alternative short-run models. One weights cells by their size; another reweights observations within each cell to give more weight to fertility outcomes with shorter recall periods (using a Bartlett kernel); another omits country, year, and age effects; and three others add country-specific linear, quadratic, and cubic time trends. Both the weighted model and the trend models deliver results very similar to those reported in Table 2. However, in the model with no country, year, or age effects, the coefficients at prime childbearing ages (20-34)—while still statistically significant—shrink by roughly one-quarter, while the coefficient in the 40-44 age group grows by the same proportion.

That the omission of country, year, and age effects modifies coefficients across the age distribution in different directions is easily reconciled by the analysis of long-run rates of change, where the results are nearly opposite the short-run estimates. As shown in the final two columns of Table 2, long-run economic growth and long-run fertility change are negatively correlated at prime ages but positively correlated at older ages. Column (5) reports the basic model without covariates, revealing these marked contrasts across the age distribution. A comparison of women in their early thirties with women in their early forties provides the starkest contrast. Among 25-29 and 30-34 year olds, a 1-point faster average annual rise in log GDP per adult is associated with a 0.43-point faster average decline in conception rates: roughly equal and opposite in sign from the short-run coefficient in column (3). Meanwhile, among 40-44 year olds the same increase in long-run growth is associated with a 0.14-point slower average decline in conception rates. This result reflects the postponement of childbearing in faster-growing economies. On net, the offsetting coefficients different ages imply a long-run TCR coefficient of -6.1, so that overall, faster long-run economic growth is associated with more rapid fertility decline. The average rate of change in log GDP per adult has a standard deviation of 1.8, so a one standard deviation increase in long-run growth is associated with fertility declining at a decadal rate of one conception for every nine women.

As with the short-run analysis, these long-run results are robust to alternative regression specifications. In column (6), the estimates remain similar after controlling for the initial level of GDP per adult as well as the average rates of change in conflict, female education, urbanization, and infant mortality. Appendix Table A3 again reports the coefficients on these covariates. Average rates of change in female education and fertility are negatively correlated, yet this relationship does not explain the long-run growth-fertility link. Further specification checks are reported in Appendix Figure A3, which plots age-group-specific coefficients from a range of alternative long-run models. Changing the minimum long-run time horizon from 20 years to 15 or 25 does not change the estimates, nor does reweighting conception rates using a Bartlett kernel as in (Appendix Figure A2). We also obtain similar results when we use the average of observed annual changes, which has the same interpretation as our main measure of the average annual change but loses information from gaps in the country-year panel.

4.4 Exploring the Procyclicality of Fertility

4.4.1 Non-Linearities

All of the preceding models assume linearity, but the theoretical framework's emphasis on liquidity constraints suggests that conception rates may respond differently to booms and busts, and these responses may be non-linear in the size of the economic fluctuation. To assess such non-linearities in the short-run results, we discretize the distribution of $100 \times \Delta \ln(GDPpa_{ct})$ into six bins (< -10, $[-10, -5)$, $[-5, 0)$, $[0, 5)$, $[5, 10)$, and ≥ 10) and then run a semi-parametric version of equation (6) that replaces the continuous variable g_{ct} with bin indicators. We prefer this approach to non-parametric regression because it allows us to control for country, year, and age effects while still measuring the change in log GDP per adult in its original units. Results are summarized in Figure 2, with Panel A plotting the age-group-specific coefficients and Panel B plotting the implied TCR coefficients with 90 and 95 percent confidence intervals. Confidence intervals for the age-group-specific coefficients are displayed in Appendix Figure A4.

In Figure 2, a clear asymmetry emerges, with conceptions falling sharply in deep recessions but not rising in pronounced expansions. Relative to the base category, moderate economic growth of 0-5 log points, a recession of more than 10 log points decreases the total conception rate by 171 per 1000 women: nearly one-fifth of a child per women. Looking across age groups, 25-29 year olds are hardest hit, with conceptions declining 16 per 1000 during these extreme recessions. As shown in the histogram at the bottom of Panel B, recessions of this magnitude are rare but not unprecedented, with 3 percent of the sample (1,644 cells) in this category. Meanwhile, during episodes of positive growth, changes in conception rates do not depend on the extent of growth.

4.4.2 Lags

The theoretical framework also makes clear that women may offset past fertility adjustments, suggesting that we should include lagged growth rates as additional covariates in equation (6). Figure 3 extends the model with four lags; alternative lag lengths produced similar results.[30] Panel A reports results for each age group, showing pronounced offset behavior in the year following a growth fluctuation. In response to a 1-point positive fluctuation, conceptions among 25-29 year olds rise 0.53 per 1000 in the same year but fall 0.40 per 1000 in the following year. Both coefficients are statistically significant, as shown in Appendix Figure A5, which draws the impulse response functions with confidence intervals. Most other age groups display the same pattern, although the confidence intervals are sometimes too wide for a definitive interpretation, especially at longer lags. One notable exception is the impulse response function for 30-34 year olds, which displays the weakest offset behavior. This pattern will be relevant for the cohort results in Section 5.

[30] To focus on the growth fluctuations that identified the results in Table 2, we only use the sample that has both lagged growth rates and lagged conception rates, which is 86 percent of the sample missing lagged conception rates.

Aggregation across age groups allows us to characterize the impulse response function with less noise. Panel B of Figure 3 reports the implied TCR coefficients for the contemporaneous and lagged effects. In response to a 1-point growth fluctuation, TCR rises 10.8 at first but then falls 7.6 in the following year. These coefficients reflect the partial effect of past economic fluctuations on the current change in the conception rate. The running sum of the coefficients, plotted in black, reveals that the effect of a fluctuation dissipates over time but does not disappear. In particular, the cumulative effect of a fluctuation on TCR shrinks to 3 one year after the fluctuation but settles at a significant 6-8 children per 1000 women thereafter. Some but not all of the short-run fertility response to economic fluctuations is offset through subsequent adjustments to childbearing. Importantly, because we study the year-to-year change in the conception rate at a fixed age, rather than the change for a fixed birth cohort, this exercise does not map exactly onto the evolution of fertility choices over time for a particular woman. The cohort analysis in Section 5 will address this issue.

4.4.3 Heterogeneity

Who responds most to growth fluctuations? Our rich dataset also allows us to explore which individual and aggregate characteristics are linked with stronger procyclicality. To shed light on heterogeneity within countries, we study how the average characteristics of mothers change over the business cycle. An alternative approach would categorize women by characteristics and then compute separate conception rates by category, but this approach considerably reduces cell sizes and also alters the country composition of our sample.[31] By looking at the average characteristics of mothers, we avoid these problems. To further minimize changes in sample composition from cells with no births, we run the analysis at the country-year level, for the same countries and years in the main sample.

We focus on four characteristics that may be especially relevant to understanding the procyclicality of conceptions: age, education, urban residence, and ever-marriage. Education and urban residence correspond to the time of the survey—at the end of the birth history—while age and marriage are contemporaneous to the year of conception risk. This approach to studying heterogeneity requires us to control carefully for the age structure and average characteristics of all women in the country-year cell. As such, we modify equation (6) to include additional covariates: changes in the fraction of all women in each single-year age group, as well as changes in average years of education,

[31] Some countries have few women with no education, for example; others have few to none with secondary education.

percent urban, and percent ever-married among all women. The coefficient on g_{ca} now captures how growth fluctuations influence the composition of conceiving mothers, over and above any association of growth with the composition of women at risk for conceiving. Before looking at maternal characteristics, however, we first estimate the modified model using the change in the conception rate (in the country-year cell) as the dependent variable. Reported in column (1) of Table 3, the coefficient is a statistically significant 0.24—which implies a TCR coefficient of 7.3, roughly four-fifths of the short-run TCR coefficients estimated more flexibly in Table 2. Thus, we obtain similar results despite aggregating age groups.

The next four columns of Table 3 present the results for maternal characteristics, with a single noteworthy result: a significant positive relationship between growth fluctuations and the average education of mothers. A log point positive growth fluctuation is associated with a decrease in average education of 0.002 years. To put this in perspective, if we replace the growth rate with an indicator for recession ($g_{ct} < 0$), we obtain a coefficient of 0.02, suggesting that the average education of mothers increases by 0.02 years (0.5 percent of the average education level and 25 percent of its annual rate of change) during a downturn. This finding implies that less-educated women are more responsive to growth fluctuations: consistent with a role for credit constraints and unsupportive of the idea that only more-educated women can control their fertility.[32]

Other average characteristics of mothers do not vary significantly with growth fluctuations. The most noteworthy of these null results concerns marriage, given the historical importance of marriage as a method of fertility limitation (Hajnal 1965; Wrigley 1981). In Appendix Table A3, we investigate this margin further by analyzing marriage rates and conception rates inside and outside marriage.[33] Conceptions are 2.4 times more procyclical after marriage than before marriage, but they are also 2.8 times more likely, which explains the stability of the share of mothers ever married. Additionally, neither the rate of new marriages among all women nor the hazard of marriage among never-married women varies significantly with growth fluctuations.

In the final column of Table 3, we use an average characteristic of children, rather than mothers, as the outcome: the fraction male. Growth fluctuations do not affect the sex composition of births, providing suggestive evidence that our fertility results, despite their reliance only on live births,

[32]The results in Table 3 do not rule out countercyclical fertility among more-educated women. But if we compute separate conception rates for more- and less-educated women, we find procyclicality in both groups, with a larger magnitude among the less-educated.

[33]We again use country-year cells, controlling flexibly for changes in the age structure.

reflect changes in conceptions rather than fetal deaths. Trivers and Willard (1973) proposed that natural selection leads mothers in worse condition to produce more female offspring, a theory that receives support in data on the sex ratio at birth in humans (Almond and Edlund 2007; Almond et al. 2016). If live births responded to growth fluctuations because of fetal death rather than fertility behavior, then this theory would predict procyclicality in the fraction male.

While the analyses in Table 3 speak to who responds most to growth fluctuations within a country, heterogeneity across countries is also of interest. Appendix Tables A4 and A5 investigate heterogeneity by aggregate characteristics and region, respectively. In Appendix Table A4, the short-run TCR coefficient does not vary significantly with the lagged levels of GDP per adult, contraceptive prevalence, average education, or urbanization, nor with the female labor force share.[34] These null results are somewhat surprising. Many results so far—procyclicality, asymmetry to booms and busts, heterogeneity by education—are consistent with a role for liquidity constraints. But conceptions are not more procyclical in poorer countries, nor in lower education countries. The theory would have also predicted more procyclicality in countries where women's work is rare, so that substitution effects are small. But the data show no heterogeneity by the extent of women's work. At the same time, Appendix Table A5 reveals marginally significant regional variation, with countries in Africa, Latin America, and the Caribbean exhibiting stronger procyclicality than those in Asia. In fact, conception rates appear to be entirely unresponsive to economic fluctuations in South and Southeast Asia. Given the null result for heterogeneity by aggregate socioeconomic characteristics, this regional heterogeneity is puzzling.

4.4.4 Comparison with Other Datasets

The suggestive evidence on regional variation raises the question of how the short-run fertility response to economic fluctuations in WFS/DHS countries compares with that in developed countries, which have been the focus of the literature on fertility and the business cycle (Sobotka, Skirbekk, and Philipov 2011). Cross-national data on conception rates by year in developed countries are not readily available, so we rely on birth rates to explore this issue. Specifically, we analyze annual age-specific birth rates from our WFS/DHS microdata and from the Human Fertility Database (HFD),

[34] To maximize interpretability, Appendix Table A4 splits the sample above and below the median of each aggregate characteristic. If we instead interact each characteristic with the change in log GDP per adult, we still find no significant heterogeneity. We measure the female labor force share in a single year, 1990, because panel data are limited for many countries. The WDI has more complete panel data after 1990, but country effects account for over 95% of the variation.

a compilation of natality data from populations with high-quality vital registration systems, mostly in developed countries. Using equation (6), we relate changes in these birth rates to the weighted average of current and lagged changes in log GDP per adult, assigning weight $\frac{1}{4}$ to the current change and $\frac{3}{4}$ to the lagged change to roughly match the conception period for the current year's births.

Table 4 presents the results, revealing comparable birth rate procyclicality. In the WFS/DHS sample, the results closely tracked the conception rate results in Table 2, with column (2) reporting a total fertility rate coefficient (which is analogous to the total conception rate coefficient, but for the birth rate) of 8.0. In the HFD sample, where the average population is 20 times richer and $\frac{2}{3}$ less fertile than the WFS/DHS sample, this coefficient is slightly smaller, at 5.5 (column [5]), although the difference between samples is not significant. In absolute terms, fertility is similarly procyclical in developing and developed countries, although this result also implies stronger proportional procyclicality in developed countries.

For an additional point of comparison, we rerun this analysis using total fertility rates (instead of age-specific birth rates) from the World Development Indicators, a popular dataset in cross-country research. For the developed country sample, estimates from the WDI are similar to those from the HFD, with column (6) showing a total fertility rate coefficient of 6.4. However, for the developing country sample, the WDI delivers estimates in column (3) that are insignificant and close to zero, likely because the WDI's fertility data are heavily smoothed for countries with low-quality vital registration systems. Researchers using this popular cross-country dataset would have incorrectly concluded that fertility is far more procyclical in richer, lower-fertility countries.

4.5 Exploring the Long-Run Link Between Growth and Fertility Decline

4.5.1 Non-Linearities

While theory gives no clear prediction regarding non-linearities in the long-run relationship, more flexible estimation may yield interesting results here as well. Because equation (7) includes no country or year effects, and the age effects are of minimal importance, we rely on a bivariate regression smoother to characterize non-linearities in the long run.[35] Figure 4 displays local linear regressions of the average annual rate of change in the conception rate on the average annual rate of economic growth, using a bandwidth of 2 log points. Panel A graphs the regression function for each age

[35] Estimation of equation (7) without age effects leads to extremely similar coefficients and significance levels.

group, with the domain running from the 5^{th} to the 95^{th} percentiles of that age group's distribution of average growth. To avoid clutter, we relegate confidence intervals to Appendix Figure A6.

Relative to the coefficient estimates in Table 2, the curves in Figure 4, Panel A, provide three new insights. First, unlike the short-run results, the long-run regression functions do not deviate substantially from linearity. Second, for all age groups, conception rates trended downward regardless of long-run growth rates; the estimated regression functions are uniformly negative. Third, the positive long-run coefficient for the 40-44 age group does not reflects an increase in conception rates in faster-growing countries, but instead a slower decrease.

To generate the implied regression function for the TCR, Panel B aggregates these results over the intersection of the age-group-specific domains. The dotted and dashed lines represent 90 and 95 percent confidence intervals, which we block-bootstrapped at the country level. Consistent with the patterns in Panel A, the average rate of fertility change declines linearly in the average growth rate. At the 10^{th} percentile of the average growth rate—where the economy shrank by 1.6 percent per year—the TCR declined at an annual rate of 60 conceptions per 1000 women. Meanwhile, at the 90^{th} percentile—where the economy grew by 3.2 percent per year—the TCR declined at a 33 percent faster rate: 80 conceptions per 1000 women annually.

4.5.2 Mechanisms

Although controlling for trends in conflict, female education, urbanization, and infant mortality did not alter results in Table 2, other relevant covariates were omitted because they are not available for all country-years. For the long-run analysis, however, yearly measurements are less important than for the short-run analysis. Even with gaps in the data, one can compute informative estimates of the average annual rate of change in a covariate. In Table 5, we control for the average annual rate of change in each of four covariates that are not available for the whole sample but may shed light on mechanisms: secondary school enrollment, female labor force participation, the sectoral composition of value added, and the extent of democracy. For each of these covariates, we report two regressions: one including the average annual rate of change in the covariate, one omitting it but using the restricted sample with a non-missing average annual rate of change. For reference, column (1) repeats the full-sample estimates from specification (7), first reported in Table 2, column (5). We compute the average annual rate of change using the methods described in Section 3. In estimating this rate of change for a country-year cell, we only use country-year-age cells that have data on both the conception rate and the covariate.

Our theoretical framework emphasizes the hypothesis by Galor and Weil (2000) that long-run growth raises the return to human capital investment, which increases investment per child and therefore the marginal cost of children. A test of this theory is difficulty because it would ideally use data on schooling (or returns to schooling) 1-2 decades after the fertility outcome. As an imperfect substitute, we rely on school enrollment at the time of the fertility outcome, under the assumption that parents' expectations regarding their children's returns are heavily influenced by the current desirability of schooling. In the WDI, gross enrollment ratios are available from many more countries in our sample than net enrollment ratios, so we use the former. A well-known problem with gross enrollment ratios is that they can by biased by grade repetition, and indeed, more than one-third of the country-years in our dataset have primary school ratios in excess of 100. We therefore rely on secondary school enrollment ratios, which never exceed 100 in our sample. The average rate of change in secondary enrollment is also closely related to the average rate of economic growth in our sample; a 1-log point higher average rate of growth predicts that secondary enrollment rises 0.1 percentage points faster per year. When we control for this covariate in column (2), the coefficients on average growth rise substantially. Because all but one of these coefficients are negative, they are attenuated by the inclusion of the rate of change in secondary enrollment; the TCR coefficient shrinks by roughly half. Therefore, trends in contemporaneous secondary school enrollment can partly explain our long-run results. Importantly, contemporaneous enrollment reflects the desirability of schooling, *not* the education of mothers. The results in Table 2 already established that maternal education does not explain the results.

A separate theory posits that the technological progress associated with economic growth raises the returns to women's work (Galor and Weil 1996), increasing the costs of childbearing and therefore pushing fertility downward. To examine this possibility, we first use the labor force participation rate of women over age 15, assembled by Olivetti (2014) from ILO databases. Columns (3)-(4) show no meaningful change in the long-run results from the inclusion of the average rate of change in this covariate. One complication is that the female labor force participation rate does exactly reflect the opportunity cost of women's time. Goldin (1995) points out that female labor force participation is high early in the development process, but wages are low and work is compatible with childcare: for example, work on the family farm close to home. A rising opportunity cost of children may be better reflected in the size of the service sector, which employs women outside the home at higher wages, and may therefore create more of a work-fertility tradeoff. Panel data on the sectoral composition of the labor force are extremely sparse, so we instead use data on the sectoral composition of value added (agriculture, manufacturing, services, and other). Here again, in columns (5)-(6), controlling for trends in these sectoral composition variables fails to change the long-run results. Overall, our tests fail to find evidence that women's work plays an important role in our long-run results. Data on wages or the sectoral composition of the labor force would improve these tests but are unavailable for most of our sample.

As a final exploration into the mechanisms driving the long-run results, we control for trends in the Polity IV score, a measure of democratic institutions. Democracy does not play a key role in economic theories of the fertility transition, but demographer Dyson (2013) suggests a possible link. Whatever the link, columns (7)-(8) reveal that trends in democratization do not explain the association between long-run growth and long-run fertility change.

4.5.3 Comparison with Other Datasets

Just as Table 4 compared the short-run results for developing-country with those for developed countries and alternative fertility data sources, Table 6 carries out this exercise for the long-run analyses. We again use vital statistics data from the HFD as our main source for developed countries, but for alternative fertility data, we draw on age-specific fertility rates from the UN instead of total fertility rates from the WDI. Relative to the WDI, the UN data have the advantage of being disaggregated into five-year age groups, but they are only available in five-year intervals, making them appropriate only for studying longer-run change, not annual fluctuations. Because interpolation and demographic modeling raise fewer problems for analyses of longer-run change, we anticipate that the UN data will generate results broadly similar to those from the WFS/DHS. However, while this similarity is likely to hold for the total fertility rate, the results may diverge for specific age groups. Some of the demographic modeling techniques employed by the UN (e.g., the Gompertz relational hazard model) are known to perform poorly at the extremes of the reproductive lifespan, where they tend to impose a rate and direction of fertility change similar to those occurring at prime childbearing ages. For all of these datasets, we compute average annual rates of change in log GDP per adult and the age-specific fertility rate using the methods described in Section 4.2. We then regress average growth on average fertility change for each age group.

In another stark departure from the short-run results, Table 6 reveals the long-run patterns are quite different for developing and developed countries. Despite using the birth rate instead of the conception rate, the long-run coefficients for the WFS/DHS dataset are similar in magnitude and sign to the main results in Table 2: negative at prime childbearing ages and positive toward the end of the reproductive lifespan, with a significantly negative implied coefficient for the total fertility rate. The HFD data show no such pattern. One age-group-specific coefficient is significantly negative, but the others are either close to zero or positive, and the implied coefficient for the total fertility rate is small and insignificant. Faster long-run growth does not appear to be associated with more fertility decline in high-income, low-fertility populations, suggesting that the fertility-reducing substitution effects of long-run growth are stronger during the development process.

Also in contrast to the short-run results, the standard cross-national database produces results that are not wildly incorrect. For developing countries, the implied total fertility rate coefficient is significantly negative; for developed countries, it is small and insignificant. However, consistent with bias from demographic modeling, the UN data perform poorly at the end of the reproductive lifespan. For developing countries, the 40-44 age group has a negative coefficient, perhaps because model assumptions forced fertility to trend similarly at younger and older ages. If we had used standard cross-national data instead of survey data in our main analysis, we still would have missed an important nuance in the relationship between growth and fertility change in the long run.

5 Analysis of Cohort Fertility

The period fertility analysis clarified the distinctions between short- and long-run relationships at the population level, but shed light on fertility behavior over the lifecycles of individual women. If women fully offset the short-run effects before the end of childbearing, then the observed procyclicality will not affect lifetime fertility; growth fluctuations will alter the tempo but not the quantum of fertility. As noted in the theoretical framework, full offset is more likely for fluctuations early in the reproductive lifecycle; women approaching menopause may not have time to make up for lost childbearing opportunities. To investigate these issues, this section changes the unit of analysis from the country-year-age cell to the country-cohort cell, defined as all women born in the same year in the same country. It then relates a cohort's completed fertility rate with its experience of economic growth over the reproductive lifecycle.

5.1 Methods

We regress the cohort's completed fertility rate on average economic conditions experienced in each age interval A from $[15, 19)$ to $[40, 44)$, adjusting for the cohort's average age when surveyed $\overline{a}_{cj}$, a country fixed effect λ_c, and a cohort fixed effect δ_j:

$$CFR_{cj} = \sum_{A} \beta^A \overline{g}^A_{cj} + \theta \overline{a}_{cj} + \lambda_c + \delta_j + \varepsilon_{cj} \tag{9}$$

where $\overline{g}^A_{cj}$ is the average annual change in log GDP per adult over age interval A, measured in log points. Together, the β^A coefficients capture how completed fertility responds to within-country, within-cohort differences in economic growth experienced over the lifecycle.

The country fixed effect λ_c absorbs cross-country variation in long-run economic growth, so that the β^A coefficients reflect the persistent influence of the short-run fluctuations. Importantly, however, the underlying variation is not exactly the same as that in the short-run period analysis. On the one hand, five year growth may be more similar to long-run growth; on the other, it may reflect deeper business cycle variation with greater liquidity effects. Nevertheless, aggregated into five year age intervals in this way, the model delivers more precise and concise estimates.

Without the country fixed effect, the β^A coefficients reflect cross-country variation as well. The Galor and Weil (2000) assumptions on the human capital production function predict that faster growing countries will have a lower level of fertility, so we also run equation (9) without λ_c. However, we acknowledge that cross-country variation in the level fertility raises many concerns about omitted variables, so we place less emphasis on this shortened regression specification.

5.2 Results

Table 7 reports how a cohort's completed fertility rate relates to its experience of economic growth over the reproductive lifecycle. The first three columns use our primary measure of completed fertility, the number of children ever born per 1000 women. Columns (1) omits the country fixed effect from equation (9) and finds that economic growth is negatively associated with the completed fertility rate across the age distribution. This result is consistent with the Galor and Weil (2000) prediction regarding the long-run growth rate and the level of fertility.

When we include the country fixed effect in columns (2) and (3) to isolate the effects of fluctuations around the long-run growth rate, the estimated coefficients change dramatically. Up to age 30, fluctuations have no relation with lifetime fertility, consistent with full offset of short-run responses. Offset opportunities appear to diminish thereafter, with the results indicating permanent effects of fluctuations in the 30s. Considering the theoretical prediction that parents will make smaller adjustments to fertility when those adjustments are more likely to be permanent, this age pattern lines up well with the rapid decline of the short-run coefficients after the late 20s. Net of the long-run growth rate, a 1 log point increase in the average annual growth rate experienced during 30-34 or 35-39 raises completed fertility by roughly 40 children per 1000 women. The estimated coefficients shrink slightly but remain statistically significant when we control for cohort characteristics like average education and share urban (column [3]) or only count children who survived until the survey date (column [4]).

Puzzlingly, however, any of these magnitudes far exceed effects that the short-run effects would imply if they were permanent. This accumulating effect appears inconsistent with the dissipation of the short-run effects in Figure 3. Research on the US has found similar patterns of short-run effects accumulating over the lifecycle (Currie and Schwandt 2014), although the key margin in that context is childlessness, which does not play an important role here.[36] Three points may help explain this puzzle. First, in the short-run model with lags, 30-34 is the age group with the weakest offset pattern. Second, fertility may respond non-linearly to a sustained and deep recession, which may be better reflected in a five-year average than a single-year growth measure. Third, the 95% confidence intervals for the short-run effects (with our without offset) and the completed fertility effects contain values consistent with each other.

An alternative approach to estimating the fertility impact of economic conditions over the lifecycle would use the level of GDP per adult, rather than its growth rate, in equation (9). In this approach, the country and cohort fixed effects absorb cross-country and cross-cohort variation in the level of development and the level of fertility, but they fail to address trends in both variables. As such,

[36] Childlessness rates are low in sample cohorts, averaging 4 percent, and are unrelated to within-country variation in cohort experiences of economic growth.

if we used the level of GDP per adult, the β^A coefficients would reflect the net influence of the short- and long-run associations, making them difficult to interpret. Nevertheless, for completeness, we present levels-on-levels regressions in the final two columns of Table 7. For comparability, we use the 5-year average of 100 times the logarithm of GDP per adult as our covariate; to match the first two columns of the table, column (5) omits the country fixed effects, while column (6) includes them. Unsurprisingly, the results are mixed. The specification without country fixed effects leads to both positive and negative significant coefficients, which has no coherent explanation. With country fixed effects, the coefficients are either close to (and not significantly different from) zero or significantly negative. Across cohorts within a country, a higher level of resources per capita in the prime childbearing ages is associated with lower completed fertility; long-run trends dominate short-run effects in the level-on-level estimations.

Conclusion

Conflicting evidence on the relationship between macroeconomic conditions and fertility has hampered the understanding of demographic change during the process of economic development. This paper contributes three main empirical results that illuminate this relationship over different time horizons in contemporary developing countries. First, conception rates are procyclical in the short run, falling during recessions. Second, prime-age conception rates decline with long-run economic growth, while older-age conception rates rise with it. Third, across birth cohorts within a country, higher economic growth late in the reproductive lifecycle predicts higher completed fertility. These results are broadly consistent with an extension of long-run growth models with endogenous fertility to include a lifecycle with liquidity constraints.

The short-run procyclicality is consistent with evidence on fertility responses to economic fluctuations both in historical, pre-industrial populations and in contemporary, industrialized populations (Lee 1997; Sobotka, Skirbekk, and Philipov 2011). Distributed lag models and cohort analyses suggest that economic fluctuations affect both the tempo (timing) and quantum (lifetime cumulation) of fertility (Bongaarts and Feeney 1998). The weight of the evidence suggests a role for liquidity constraints, but beyond this implication, the mechanism behind procyclicality in the short run is not clear. One possibility involves couples taking intentional steps to reduce conception risk during recessions, by using modern contraception or traditional birth control strategies like withdrawal, the rhythm method, or abstinence. Because we only measure conceptions that resulted in live birth, abortion may play a role. But other mechanisms beyond conscious choice may also be at work. Stress may decrease coital frequency among cohabiting couples, and migration for labor market opportunities may temporarily split couples (Timaeus and Graham 1989). Crisis-related malnutrition may also reduce fecundity (Bongaarts 1980). How to distinguish these possibilities is a fruitful direction for future research. Another promising direction concerns the reason for the apparent absence of procyclicality in South and Southeast Asia.[37]

The short-run patterns stand in stark contrast to the relationship between long-run trends in income per head and fertility, which is negative on average but heterogeneous across age groups. The main takeaway is that some force that accompanies long-run economic growth leads to faster declines in childbearing, as reflected in the negative long-run coefficient for the total conception rate, and also to delays in childbearing, as reflected in the positive long-run coefficient for conceptions among 40-44 years olds. Evidence suggests that this force is related to rising secondary school enrollment, but *not* declining child mortality, rising adult female education or labor force participation, structural transformation, or democratization. Theories positing that long-run economic growth raises the return to human capital investment (Galor and Weil 2000)—either in children, explaining fertility decline, or in women as adults (including on-the-job training), explaining fertility delay—may therefore go a long way in explaining the long-run results.[38]

While our results help clarify divergent results on the relationship between aggregate income growth and fertility change, they raise interesting questions about mechanisms and about how fertility's relation to economic growth varies with the underlying source of that growth. They also highlight the importance of careful measurement, showing how one can use large amounts of retrospective survey data to improve on standard cross-national datasets.

[37]This result is especially puzzling because analyses of India and Indonesia find within-country patterns that are similar to our main findings: procyclicality in fertility or contraceptive use (McKelvey, Thomas, and Frankenberg 2012; Bhalotra and Rocha 2016) amidst a negative association between growth and fertility change in the long run (Gertler and Molyneaux 1994; Foster and Rosenzweig 2006) .

[38]Reversals in intergenerational wealth flows (Caldwell 1982), individualization or secularization (Lesthaeghe 1983) may also play a role, although some of these processes are consistent with rising child investment returns.

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Table 1: Summary Statistics

	Period analysis		Cohort
	Full sample	20-year sample	analysis
	(1)	(2)	(3)
Fertility rates (per 1000)			
Age-specific conception rate	191 (95)	199 (91)	
Completed fertility rate, ever-born			5951 (1284)
Completed fertility rate, surviving			4862 (870)
Macroeconomic conditions			
Country GDPpa, PPP	4711 (4218)	4476 (4087)	4229 (3292)
Change in log GDPpa, log pts.	1.0 (6.1)	1.0 (5.8)	0.7 (1.6)
Cell characteristics			
Average years of education	4.6 (5.7)	4.0 (2.9)	3.6 (2.4)
Percent urban at survey	40 (21)	38 (22)	37 (22)
Number of women	2,374,019	2,279,955	242,886
Number of cells	67,050	58,988	935
Number of countries	76	65	62

Notes: Period sample consists of country-year-age cells; cohort sample consists of country-cohort cells. "Conception rate" only includes conceptions that resulted in live birth. "GDPpa" is gross domestic product per adult age 15-64. For the cohort sample, macroeconomic conditions are first averaged over each cohort's reproductive lifecycle and then summarized across cohorts.

Table 2: Economic Growth and Conception Rates in the Short- and Long-Run

	Mean of conception rate per 1000 in...		Short run regressions		Long run regressions	
	Levels	Changes	Basic	Extended	Basic	Extended
	(1)	(2)	(3)	(4)	(5)	(6)
Ages 15-19	164	-1.2	0.10	0.12	-0.08	-0.06
			[0.07]	[0.07]*	[0.07]	[0.08]
Ages 20-24	261	-1.9	0.30	0.33	-0.27	-0.27
			[0.09]***	[0.09]***	[0.13]**	[0.13]**
Ages 25-29	253	-2.4	0.56	0.59	-0.43	-0.46
			[0.14]***	[0.14]***	[0.11]***	[0.11]***
Ages 30-34	209	-2.4	0.33	0.35	-0.43	-0.46
			[0.16]**	[0.15]**	[0.14]***	[0.11]***
Ages 35-39	144	-2.6	0.22	0.21	-0.15	-0.07
			[0.19]	[0.19]	[0.14]	[0.09]
Ages 40-44	65	-2.9	0.23	0.22	0.14	0.16
			[0.20]	[0.20]	[0.05]***	[0.05]***
TCR	5483	-67	8.77	9.07	-6.14	-5.77
			[2.50]***	[2.58]***	[2.29]***	[1.80]***
# cells	58,994	56,926	56,926	56,926	1,595	1,595

Notes: Columns (1)-(4) use country-year-age cells; columns (5)-(6) use country-age cells. Columns (3)-(4) regress the annual change in the age-specific conception rate on the annual change in 100 $\times$ log GDP per adult, controlling for country, year, and single-year age effects; columns (5)-(6) regress the average annual rate of change in the age-specific conception rate on average annual rate of economic growth, controlling for single-year age effects. "Extended" models also control for the initial level of GDP per adult (PPP) and population density; and the change or trend in female education, urbanization, infant mortality, and conflict. Column (4) also an indicator for missing mortality information (3% of all cells). "Conception rate" only includes conceptions that resulted in live birth; "TCR" refers to the total conception rate per 1000; estimates equal 5 times the sum of age-group-specific estimates. Brackets contain standard errors clustered by country. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Table 3: Cyclicality in the Composition of Births

		Average characteristics of...				
		Mothers				Children
	Concep. rate	Age	Education	% urban	% ever mar.	% male
	(1)	(2)	(3)	(4)	(5)	(6)
Δ log GDPpa × 100	0.24	-0.0001	-0.0022	-0.014	-0.009	0.012
	[0.07]***	[0.001]	[0.0009]***	[0.013]	[0.014]	[0.016]
Outcome mean	201	23	3.5	35	92	51
Outcome SD	(52)	(3)	(2.5)	(19)	(10)	(3)
# cells	2831	2831	2831	2831	2831	2831

Notes: Regressions of annual changes in average characteristics on annual changes in 100 × log GDP per adult, controlling for country and year fixed effects, as well as changes in age composition, average years of education, percent urban, and percent married among all women in each cell. "Conception rate" only includes conceptions that resulted in live birth. Brackets contain standard errors clustered by country. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Table 4: Comparison of Procyclicality Results with Other Datasets

	Country-years in the WFS/DHS Mean 2005 GDPpa, PPP = 5,239 Mean GDPpa growth = 0.91			**Country-years in the HFD** Mean 2005 GDPpa, PPP = 46,993 Mean GDPpa growth = 2.41		
	WFS/DHS		WDI	HFD		WDI
	Mean rate (1)	Regression (2)	Regression (3)	Mean rate (4)	Regression (5)	Regression (6)
Ages 15-19	138	0.163		28	0.173	
		[0.078]**			[0.062]***	
Ages 20-24	258	0.145		103	0.397	
		[0.097]			[0.135]***	
Ages 25-29	261	0.475		126	0.179	
		[0.108]***			[0.086]**	
Ages 30-34	224	0.340		83	0.189	
		[0.107]***			[0.066]***	
Ages 35-39	161	0.358		35	0.141	
		[0.151]**			[0.029]***	
Ages 40-44	80	0.155		8	0.030	
		[0.277]			[0.008]***	
Total fertility rate per 1000	5601	8.19	0.28	1920	5.55	6.44
		[2.63]***	[0.23]		[1.14]***	[1.31]***
Num. of cells	57,126	55,479	2,460	23,310	23,130	760

Notes: "WFS" = World Fertility Survey; "DHS" = Demographic and Health Survey; "HFD" = Human Fertility Database; "WDI" = World Development Indicators. Coefficients from regressions of annual changes in the age-specific fertility rate on the weighted average of current and lagged annual changes in 100 × log GDP per adult, with weight 0.25 on the current change and weight 0.75 on the lagged change. In the WFS/DHS and HDI, unit of observation is a country-year-age cell, and the dependent variable is the age-specific birth rate; analyses are run by 5-year age group and include country, year, and age fixed effects. Total fertility rate estimates equal 5 times the sum of age-group-specific estimates. In the WDI, unit of observation is a country-year cell, and the dependent variable is the total fertility rate; analyses are adjusted for country and year indicators. Brackets contain standard errors clustered by country. Sample includes all WFS/DHS and HFD cells that can be matched with macroeconomic data from the Penn World Table and total fertility rate data from the WDI, excluding cells with < 30 obs and from country-age combinations spanning < 20 yrs. WFS/DHS countries are listed in Table A1; HFD countries include Austria, Belarus, Bulgaria, Canada, Czech Republic, Estonia, Finland, France, Germany, Hungary, Iceland, Japan, Lithuania, Netherlands, Norway, Portugal, Russia, Slovakia, Slovenia, Sweden, Switzerland, Ukraine, United Kingdom, and the United States. We omit Japanese data for 1966, when birth rates dropped 25% due to superstition surrounding the year of the fire horse. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Table 5: Alternative Long-Run Covariates

	Secondary school gross enrollment rate		Sectoral composition of value added		Female lab. force participation		Sample with POLITY IV Score	
	(1)	(2)	(3)	(4)	(3)	(4)	(5)	(6)
Ages 15-19	-0.06	-0.02	-0.11	-0.11	-0.06	-0.05	-0.14	-0.09
	[0.10]	[0.11]	[0.10]	[0.09]	[0.08]	[0.10]	[0.07]*	[0.07]
Ages 20-24	-0.43	-0.36	-0.27	-0.27	-0.33	-0.37	-0.38	-0.38
	[0.16]**	[0.17]**	[0.14]*	[0.15]*	[0.13]**	[0.17]**	[0.11]***	[0.11]***
Ages 25-29	-0.42	-0.33	-0.42	-0.44	-0.46	-0.52	-0.48	-0.49
	[0.14]***	[0.17]	[0.15]***	[0.15]***	[0.12]***	[0.13]***	[0.11]***	[0.13]***
Ages 30-34	-0.29	-0.12	-0.42	-0.45	-0.35	-0.42	-0.42	-0.41
	[0.14]**	[0.15]	[0.17]**	[0.17]***	[0.15]**	[0.20]**	[0.14]***	[0.16]***
Ages 35-39	-0.08	0.06	-0.39	-0.41	-0.08	-0.05	-0.12	-0.08
	[0.15]	[0.16]	[0.16]**	[0.16]***	[0.15]	[0.19]	[0.13]	[0.13]
Ages 40-44	0.20	0.28	0.07	0.05	0.14	0.17	0.14	0.14
	[0.07]***	[0.07]***	[0.08]	[0.09]	[0.06]**	[0.08]**	[0.05]**	[0.05]**
TCR	-5.35	-2.42	-7.74	-8.12	-5.69	-6.16	-7.00	-6.57
	[2.92]*	[3.09]	[3.00]***	[2.96]***	[2.50]**	[3.16]*	[2.16]***	[2.30]***
Covariate?		✓		✓		✓		✓
# cells	1297	1297	1424	1424	1261	1261	1532	1532

Notes: Regressions the average annual rate of change in the age-specific conception rate on the average annual rate of economic growth, as reported in columns (5)-(6) of Table 2. Each pair of columns restricts to the subsample with non-missing information on the average annual rate of change in the specified covariate. The even-numbered columns report models that include an age-specific coefficient on the average annual rate of change in the covariate. "TCR" refers to the total conception rate per 1000; estimates equal 5 times the sum of age-group-specific estimates. Brackets contain standard errors clustered by country. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Table 6: Comparison of Long-Run Results with Other Datasets

	Country-ages in the WFS/DHS		Country-ages in the HFD	
	WFS/DHS	UN	HFD	UN
	(1)	(2)	(3)	(4)
Ages 15-19	-0.116	-0.008	0.191	0.36
	[0.069]	[0.090]	[0.160]	[0.152]**
Ages 20-24	-0.286	-0.354	-0.181	0.511
	[0.126]**	[0.109]***	[0.482]	[0.263]*
Ages 25-29	-0.476	-0.562	-0.673	-0.184
	[0.116]***	[0.148]***	[0.314]**	[0.247]
Ages 30-34	-0.423	-0.529	-0.073	-0.329
	[0.132]***	[0.158]***	[0.177]	[0.234]
Ages 35-39	-0.220	-0.332	-0.030	-0.275
	[0.142]	[0.107]***	[0.176]	[0.228]
Ages 40-44	0.152	-0.076	-0.024	-0.130
	[0.069]**	[0.066]	[0.083]	[0.121]
Total fertility	-6.84	-9.30	-3.95	-2.33
rate per 1000	[2.27]***	[2.52]***	[3.58]	[3.35]
Num. of cells	1601	317	510	96

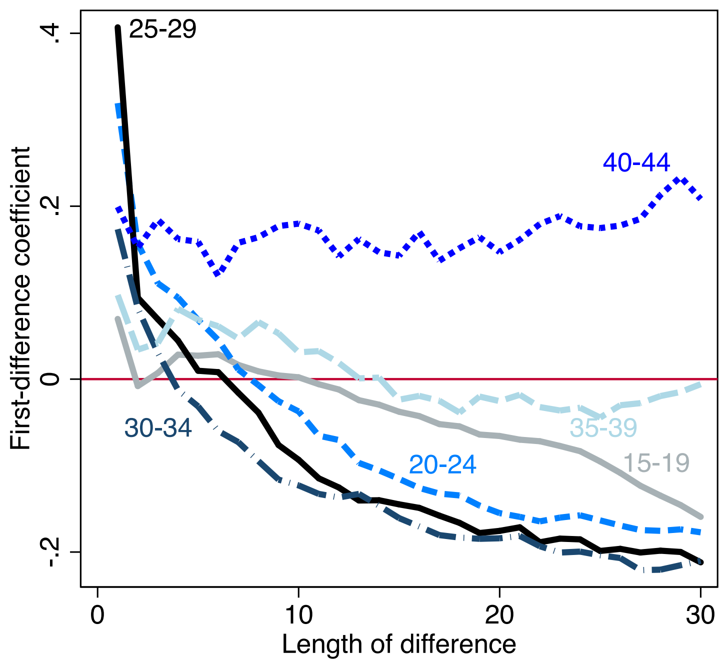
Notes: "WFS" = World Fertility Survey; "DHS" = Demographic and Health Survey; "HFD" = Human Fertility Database; "UN" = United Nations *World Population Prospects*, 2015 Revision. Coefficients from regressions of the average annual rate of change in the conception rate on the average annual rate of economic growth. The unit of observation is a country-age cell, and the dependent variable is the average annual rate of change in age-specific birth rate. Total fertility rate estimates equal 5 times the sum of age-group-specific estimates. Brackets contain standard errors clustered by country. WFS/DHS countries are listed in Table A1; HFD countries include Austria, Bulgaria, Canada, Finland, France, Germany, Hungary, Iceland, Japan, Netherlands, Norway, Portugal, Sweden, Switzerland, Taiwan, United Kingdom, and the United States. We omit Japanese data for 1966, when birth rates dropped 25% due to superstition surrounding the year of the fire horse. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Table 7: Economic Growth and Completed Fertility Across Cohorts Aged 45+

	Economic conditions measured as average...					
	Annual change in 100 × log GDPpa				100 × log GDPpa, PPP	
	(1)	(2)	(3)	(4)	(5)	(6)
Avg. economic conditions during ages...						
15-19	-55	-7	-15	-5	2.9	-0.9
	[18]***	[13]	[10]	[12]	[3.9]	[2.3]
20-24	-46	-3	-15	2	2.9	0.9
	[21]**	[14]	[12]	[14]	[4.3]	[2.8]
25-29	-66	10	1	7	-2.5	-3.6
	[17]***	[16]	[15]	[16]	[3.4]	[1.6]**
30-34	-42	38	26	29	-7.4	-4.2
	[14]***	[14]**	[12]**	[12]**	[3.8]*	[1.7]**
35-39	-46	43	35	36	6.8	1.9
	[20]***	[13]*	[12]***	[12]***	[3.7]*	[1.6]
40-44	-46	25	34	18	-11.4	-1.0
	[29]	[17]	[12]**	[15]	[3.3]***	[1.4]
Cohort avg. ed.			-226			
			[54]***			
Cohort % urban			-4.7			
			[5.3]			
Cohort FE	✓	✓	✓	✓	✓	✓
Country FE		✓	✓	✓		✓
Fertility measure	Ever-born	Ever-born	Ever-born	Surviving	Ever-born	Ever-born
Num. cells	935	935	935	935	935	935

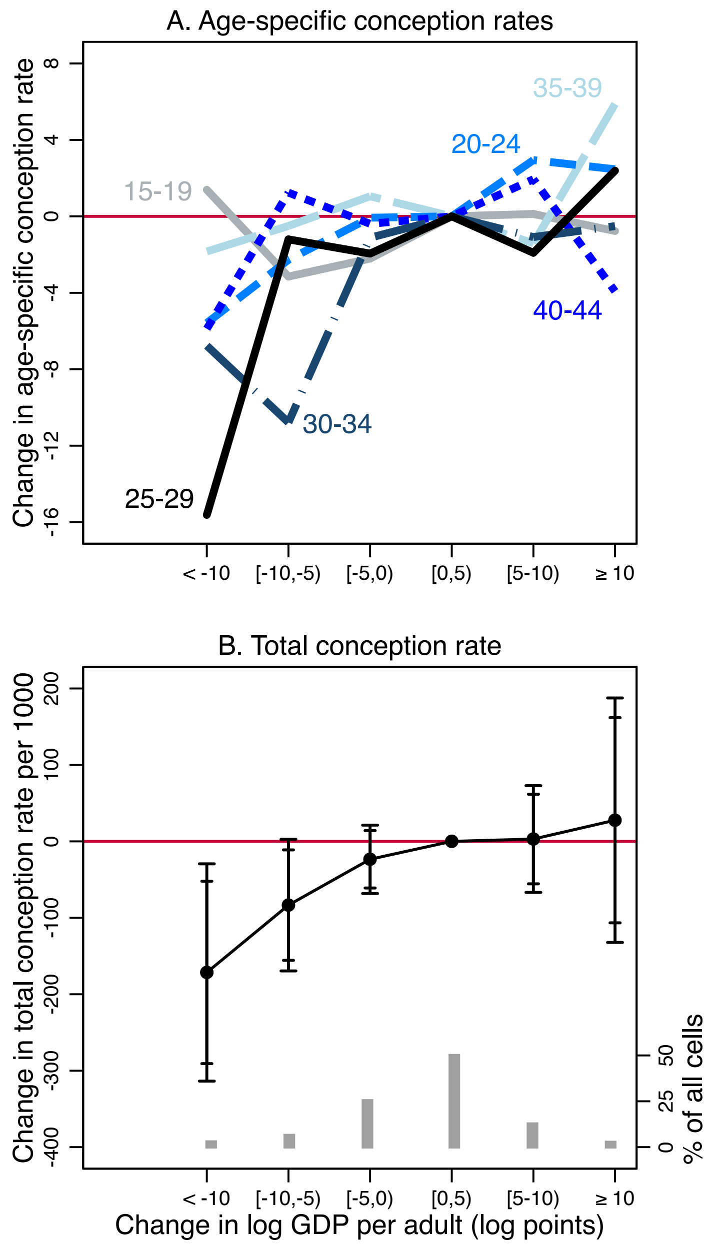
Notes: Dependent variable is the number of children per 1000 women. All regressions control for the average age (45-49) of the cohort when surveyed. Brackets contain SEs clustered by country. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Figure 1: Economic Growth and Fertility Change over Varying Time Horizons



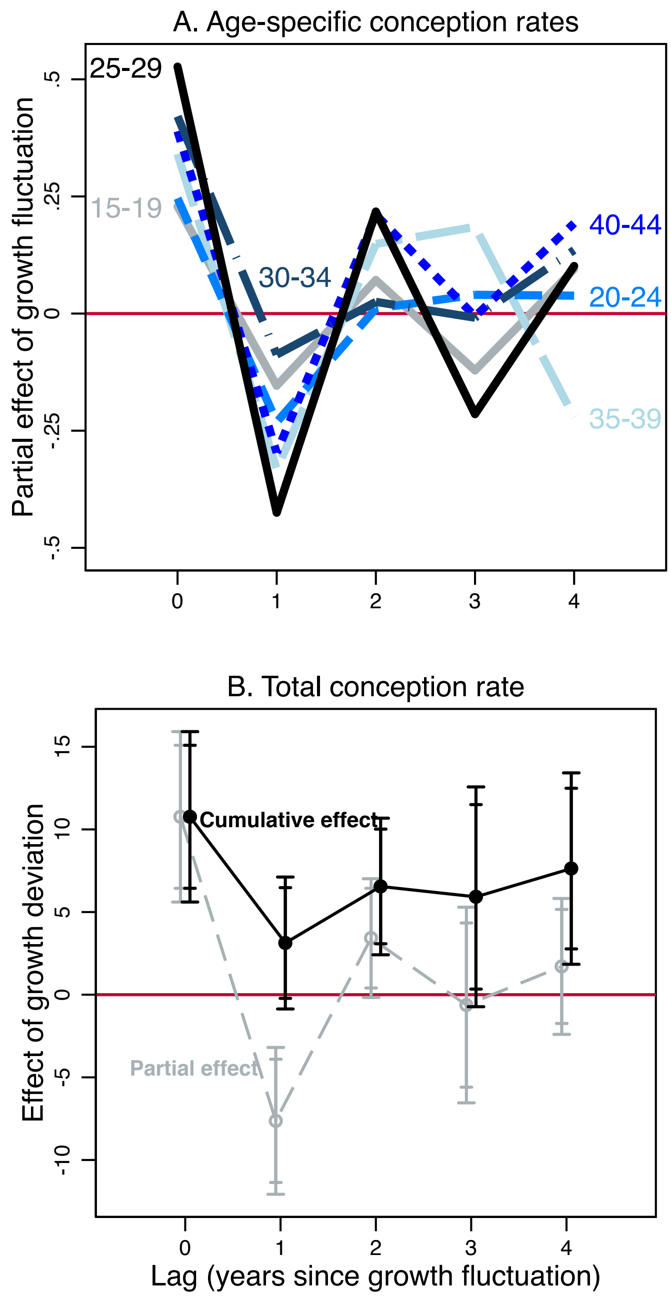
Notes: For each 5-year age group, the figure plots coefficients from regressions of the change in the conception rate from year $t - \Delta$ to year t on the change in 100 $\times$ log GDP per adult over the same period, controlling for single-year age indicators. Separate regressions were run for each integer value of Δ from 1 to 30.

Figure 2: Non-Linear Short-Run Estimates



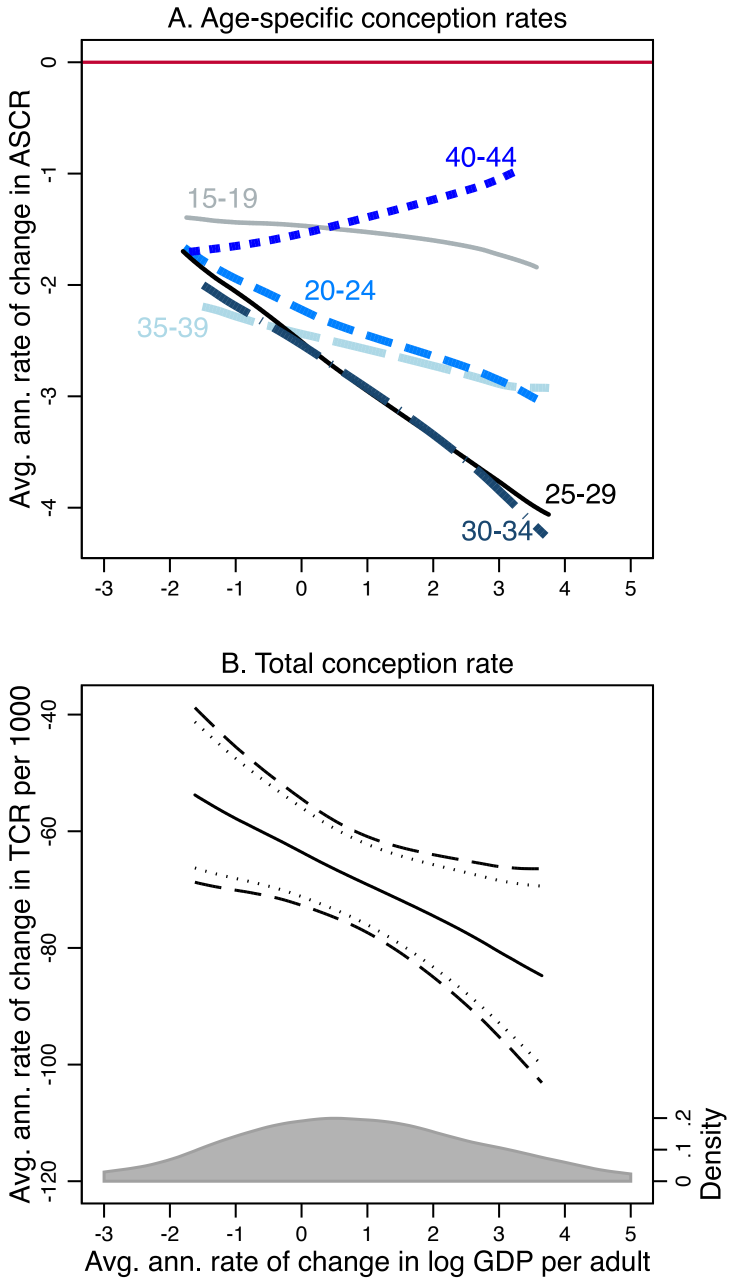
Notes: Panel A shows coefficients for each age group from regressions of annual changes in the age-specific conception rate on binned annual economic growth, controlling for country, year, and age fixed effects. Panel B shows implied estimates and CIs for the total conception rate per 1000 women, obtained by summing the age-group-specific estimates and multiplying by 5. Omitted category is [0,5). Major caps are 95% CIs; minor caps are 90% CIs. CIs clustered by country.

Figure 3: Distributed Lag Models



Notes: Panel A shows coefficients for each age group from regressions of annual changes in the age-specific conception rate on current and lagged annual changes in 100 $\times$ log GDP per adult, controlling for country, year, and age fixed effects. Panel B shows implied estimates and CIs for the total conception rate per 1000 women. Partial effects equal 5 times the sum of age-group-specific estimates; cumulative effects are the running sum of partial effects. Major caps are 95% CIs; minor caps are 90% CIs. CIs clustered by country.

Figure 4: Non-Linear Long-Run Estimates



Notes: Panel A shows local linear regression estimates for each age group, bandwidth = 2. The domain of each plot runs from the age group's 5th to 95th percentile of the average annual rate of change in log GDP per adult. Panel B shows implied regression function for the total conception rate per 1000 women, obtained by summing the age-group-specific estimates (for the domain in which they overlap) and multiplying by 5. CIs (dashed = 95%, dotted = 90%) are block bootstrapped by country. At the bottom of Panel B is a kernel density estimate for all country-age cells, bandwidth = 1.

APPENDIX

Table A1: Number of WFS/DHS Surveys per Country

Albania	1	Ghana	7	Pakistan	4
Armenia	3	Guatemala	2	Panama	1
Azerbaijan	1	Guinea	3	Paraguay	2
Bangladesh	8	Honduras	2	Peru	9
Benin	5	India	3	Philippines	5
Bolivia	5	Indonesia	8	Rwanda	5
Brazil	2	Jamaica	1	Sao Tome and Principe	1
Burkina Faso	4	Jordan	5	Senegal	8
Burundi	2	Kazakhstan	2	Sierra Leone	1
Cambodia	4	Kenya	7	South Africa	1
Cameroon	5	Korea, Rep.	1	Sri Lanka	1
Central African Republic	1	Kyrgyz Republic	2	Swaziland	1
Chad	2	Lesotho	4	Syria	1
Colombia	7	Liberia	3	Tajikistan	1
Comoros	2	Madagascar	4	Tanzania	5
Congo, Dem. Rep.	2	Malawi	4	Thailand	1
Congo, Rep.	2	Maldives	1	Togo	3
Costa Rica	1	Mali	3	Trinidad and Tobago	2
Cote d'Ivoire	4	Mauritania	1	Tunisia	2
Dominican Republic	8	Mexico	2	Turkey	4
Ecuador	2	Moldova	1	Uganda	5
Egypt	8	Morocco	4	Ukraine	1
El Salvador	1	Mozambique	3	Uzbekistan	1
Ethiopia	3	Namibia	4	Venezuela	1
Fiji	1	Nepal	5	Vietnam	1
Gabon	2	Niger	4	Zambia	5
Gambia	1	Nigeria	5	Zimbabwe	5

List includes all 255 World Fertility Surveys and Demographic and Health Surveys that could be matched with growth rates from the Penn World Table.

Table A2: Extended Model Results from Table 2

Age group:	15-19	20-24	25-29	30-34	35-39	40-44
	(1)	(2)	(3)	(4)	(5)	(6)
A. Short-run regression (Table 2, column 4)						
Δ ln(GDPpa)	0.12	0.33	0.59	0.35	0.21	0.22
× 100	[0.07]*	[0.19]***	[0.14]***	[0.16]**	[0.19]	[0.20]
Lagged ln(GDPpa)	1.33	0.55	0.93	1.40	-1.08	-0.37
× 100, PPP	[0.65]**	[0.79]	[0.78]	[0.97]	[0.94]	[1.26]
Δ avg. yrs. ed.	-9.21	-9.51	-6.58	-4.33	-2.46	-2.57
	[1.45]***	[1.81]***	[1.97]***	[2.09]**	[2.00]	[1.98]
Δ pct. urban	-0.04	0.12	0.05	-0.16	-0.28	-0.22
	[0.13]	[0.18]	[0.16]	[0.14]	[0.14]**	[0.17]
Δ conflict	-1.49	2.22	-0.13	-1.22	1.85	3.00
	[1.87]	[3.07]	[2.74]	[3.53]	[3.53]	[3.20]
Δ IMR_{t-1}	0.02	0.19	0.27	-0.04	-0.01	0.04
	[0.03]	[0.05]***	[0.08]***	[0.08]	[0.09]	[0.09]
Lagged Pop	0.001	0.005	-0.007	-0.004	0.001	0.018
Density	[0.005]	[0.007]	[0.004]*	[0.006]	[0.008]	[0.008]**
B. Long-run regression (Table 2, column 6)						
Trend ln(GDPpa)	-0.06	-0.27	-0.46	-0.46	-0.07	0.16
× 100	[0.08]	[0.13]**	[0.11]***	[0.11]***	[0.09]	[0.05]***
Initial ln(GDPpa)	-0.43	-1.34	-1.38	-1.04	-0.23	-0.12
× 100, PPP	[0.25]*	[0.28]***	[0.27]***	[0.35]***	[0.36]	[0.22]
Trend avg. yrs. ed.	-4.47	-5.46	-8.07	-10.36	-7.53	-6.57
	[3.84]	[3.86]	[3.50]**	[4.77]**	[4.38]*	[2.73]**
Trend pct. urban	-0.05	-0.11	0.00	0.19	0.32	-0.14
	[0.16]	[0.20]	[0.10]	[0.08]**	[0.07]***	[0.20]
Trend conflict	1.72	8.28	12.29	19.77	19.76	4.12
	[12.34]	[11.08]	[12.34]	[20.21]	[20.22]	[8.13]
Trend IMR	-0.07	0.00	0.10	0.05	-0.02	-0.02
	[0.13]	[0.15]	[0.15]	[0.33]	[0.23]	[0.12]
Initial Pop	-0.005	-0.005	-0.007	-0.007	-0.004	-0.0009
Density	[0.001]***	[0.002]**	[0.002]***	[0.002]***	[0.002]**	[0.0009]

Notes: Each panel represents a single regression; each column provides coefficients for a separate age group. Panel A regresses the annual change in the age-specific conception rate on the covariates shown, controlling for country, year, and single-year age effects. Panel B regresses the trend in the age-specific conception rate on the covariates shown. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Table A3: Can Marriage Explain the Procyclicality of Conceptions?

	Conception rate			Marriage	
	Overall	Pre-marital	Post-marital	Rate	Hazard
	(1)	(2)	(3)	(4)	(5)
Δ log GDPpa	0.23	0.14	0.33	-0.05	-0.13
× 100	[0.07]***	[0.09]*	[0.16]**	[0.06]	(0.12)
Outcome level mean	201	97	270	52	195
Outcome level SD	(52)	(36)	(78)	(23)	(73)
Number of cells	2831	2830	2831	2831	2830

Notes: Regressions of the changes in outcomes on annual changes in 100 × log GDP per adult, controlling for country and year fixed effects, as well as changes in the age composition of each cell. All rates are per 1000. Columns (2) and (5) have smaller sample sizes because 1 cell has no never-married women. Brackets contain standard errors clustered by country. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Table A4: Aggregate Heterogeneity in the Procyclicality of Conceptions

	Lagged GDPpa, PPP	Lagged contraceptive prevalence	Female labor force share in 1990	Lagged average years of education	Lagged share urban at survey
	(1)	(2)	(3)	(4)	(5)
Coefficient below	8.46	8.58	9.95	8.83	8.80
variable's median	[3.19]***	[3.32]***	[3.62]***	[3.42]***	[3.37]**
Coefficient above	9.68	10.22	8.33	9.24	7.86
variable's median	[2.65]***	[3.24]***	[3.15]***	[1.94]***	[1.78]***
p-value for difference	0.741	0.725	0.725	0.911	0.779
Number of cells	56,926	48,092	56,926	56,926	56,926

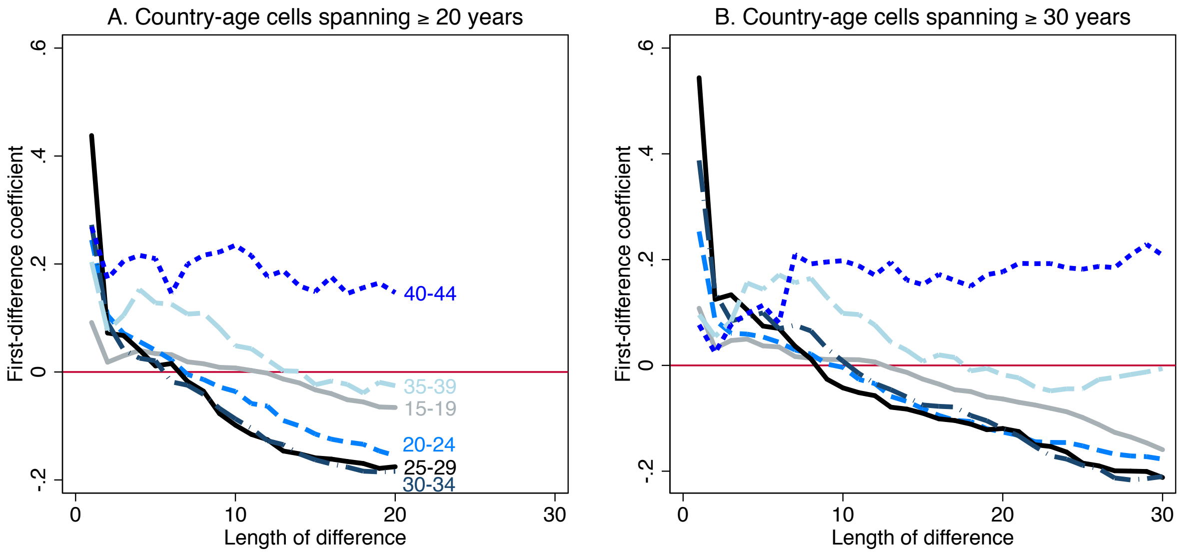
Notes: Total conception rate coefficients based on regressions of annual changes in the age-specific conception rate on annual changes in 100 $\times$ log GDP per adult, controlling for country, year, and age fixed effects. Coefficients are estimated by 5-year age group and then summed and multiplied by 5 to obtain TCR coefficient. Brackets contain standard errors clustered by country. Sample sizes vary because data on some of the aggregate variables are not available for the full sample. "GDPpa" is GDP per adult, from the Penn World Table; contraceptive prevalence is the estimated share of women of childbearing age using modern contraceptives, from the UN; female labor force share is the percent of the labor force aged 15-64 that is female, from the WDI; average years of education and share urban are estimates from WFS/DHS survey data. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Table A5: Regional Heterogeneity in Procyclicality

		p-values: coefficients equal within pair		
		Africa	C/W Asia	S/SE Asia
	(1)	(2)	(3)	(4)
Africa	9.43			
	[2.96]***			
Central/Western Asia	4.53	0.01		
	[2.96]			
South/Southeast Asia	-2.46	0.18	0.44	
	[8.43]			
Latin America/Caribbean	10.81	0.69	0.09	0.12
	[2.33]***			
p-value: all coefficients equal	0.15			
Number of cells	56,926			

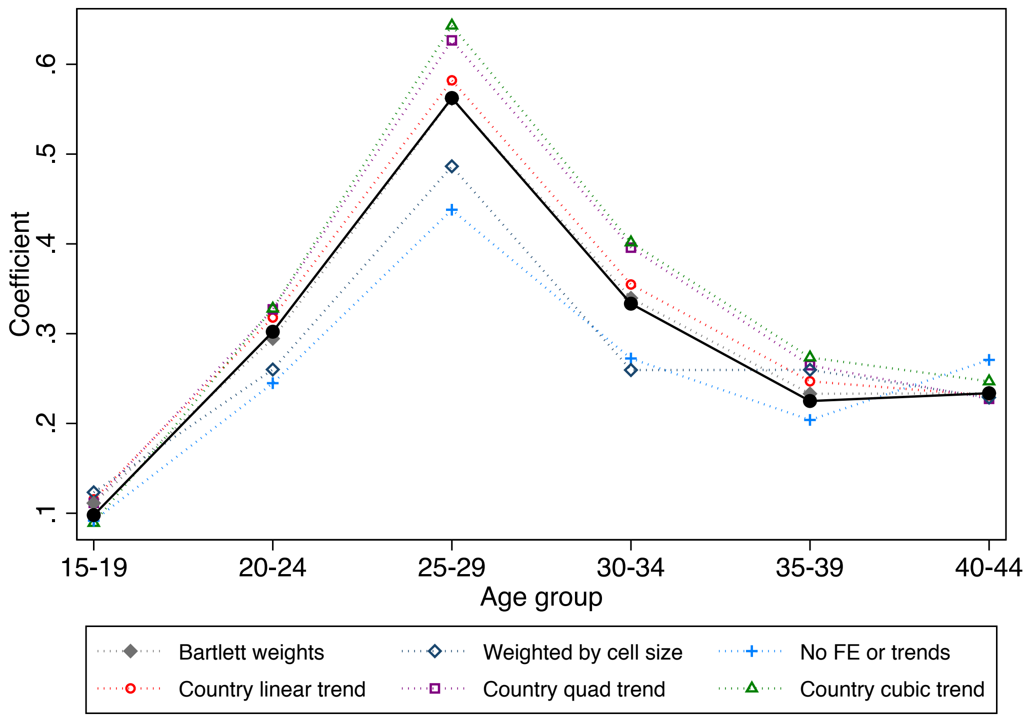
Notes: Total conception rate coefficients based on full-sample regressions of annual changes in the age-specific conception rate on annual changes in 100 $\times$ log GDP per adult interacted with region indicators, controlling for country, year, and age fixed effects. An additional (unreported) interaction term is included for the group of five countries (Albania, Fiji, Korea, Moldova, Ukraine) that did not fit into these regional classifications. Analyses are run by 5-year age group; age group associations are summed and multiplied by 5 to obtain TCR association. Brackets contain standard errors clustered by country. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Figure A1: First-Difference Models with Varying Time Horizons, Constant Samples



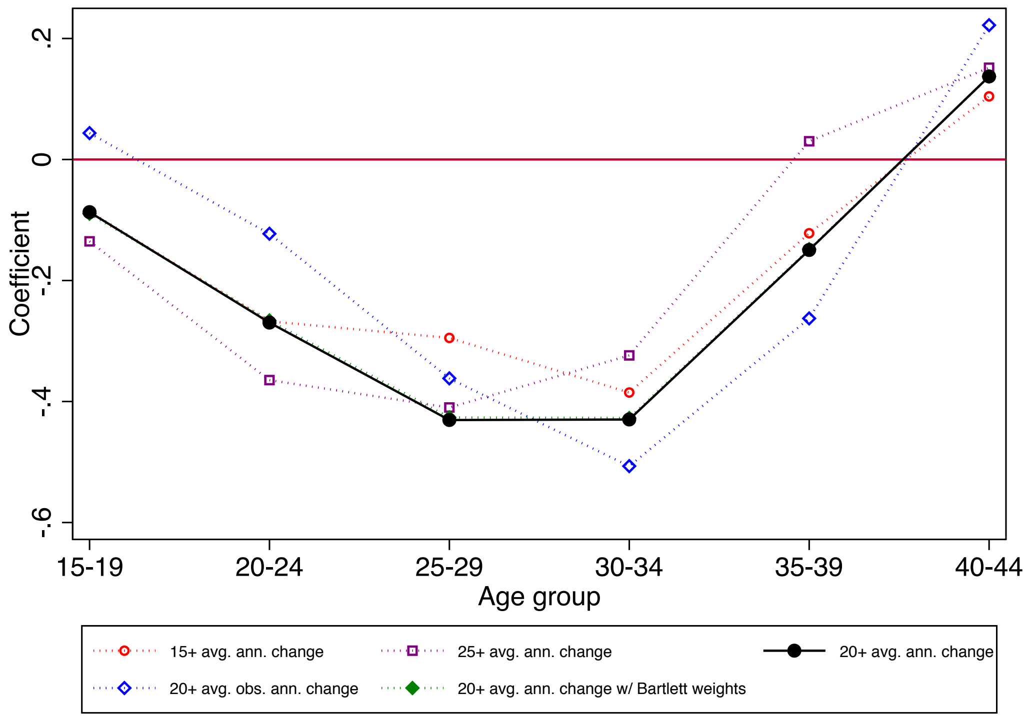
Notes: Reproduces Figure 1 using samples that do not change for different time horizons. For each 5-year age group, each panel plots coefficients from regressions of the change in the conception rate from year $t - \Delta$ to year t on the change in 100 $\times$ log GDP per adult over the same period, controlling for single-year age indicators. Separate regressions were run for each integer value of Δ from 1 to 20 (Panel A) and 1 to 30 (Panel B).

Figure A2: Alternative Short-Run Models



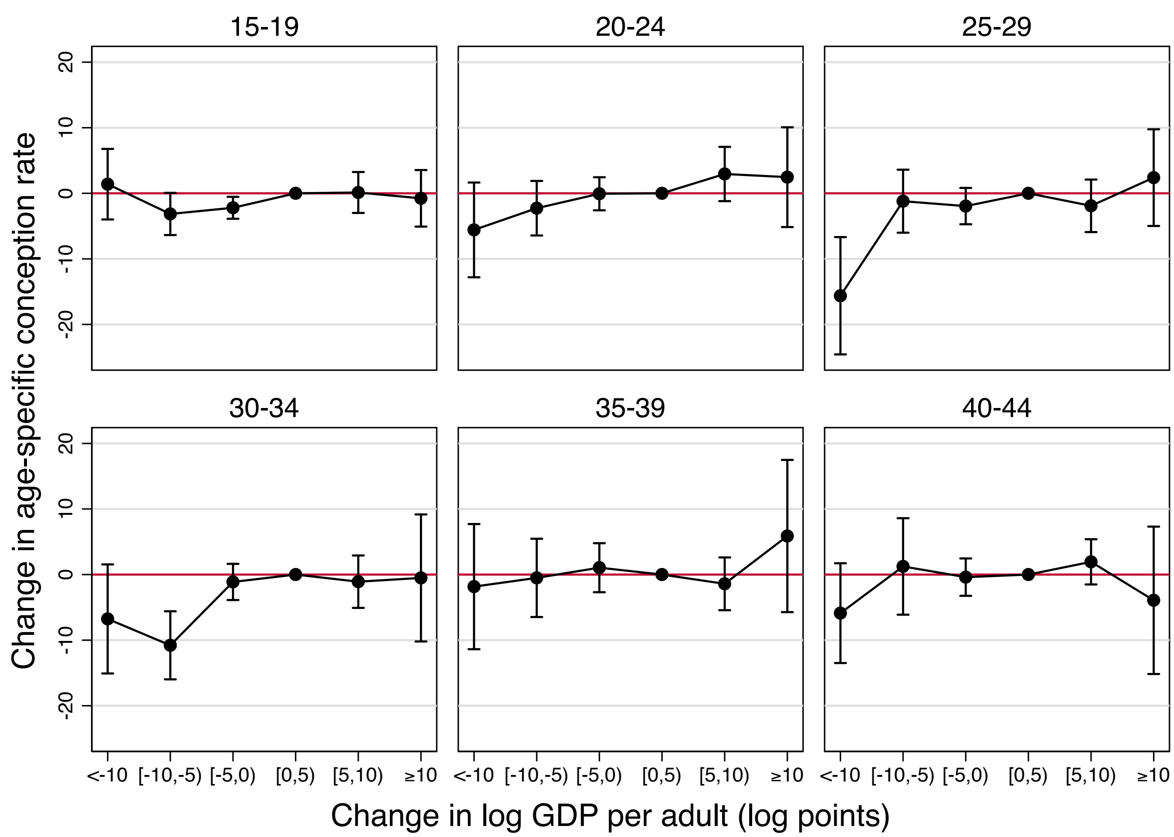
Notes: Age-group-specific coefficients from regressions of the change in the conception rate on the change in log GDP per adult. "Baseline" is the short-run model reported in Table 2; "Bartlett weights" uses a Bartlett kernel to downweight longer recall periods; trend models add country-specific polynomials in time to the baseline model.

Figure A3: Alternative Long-Run Models



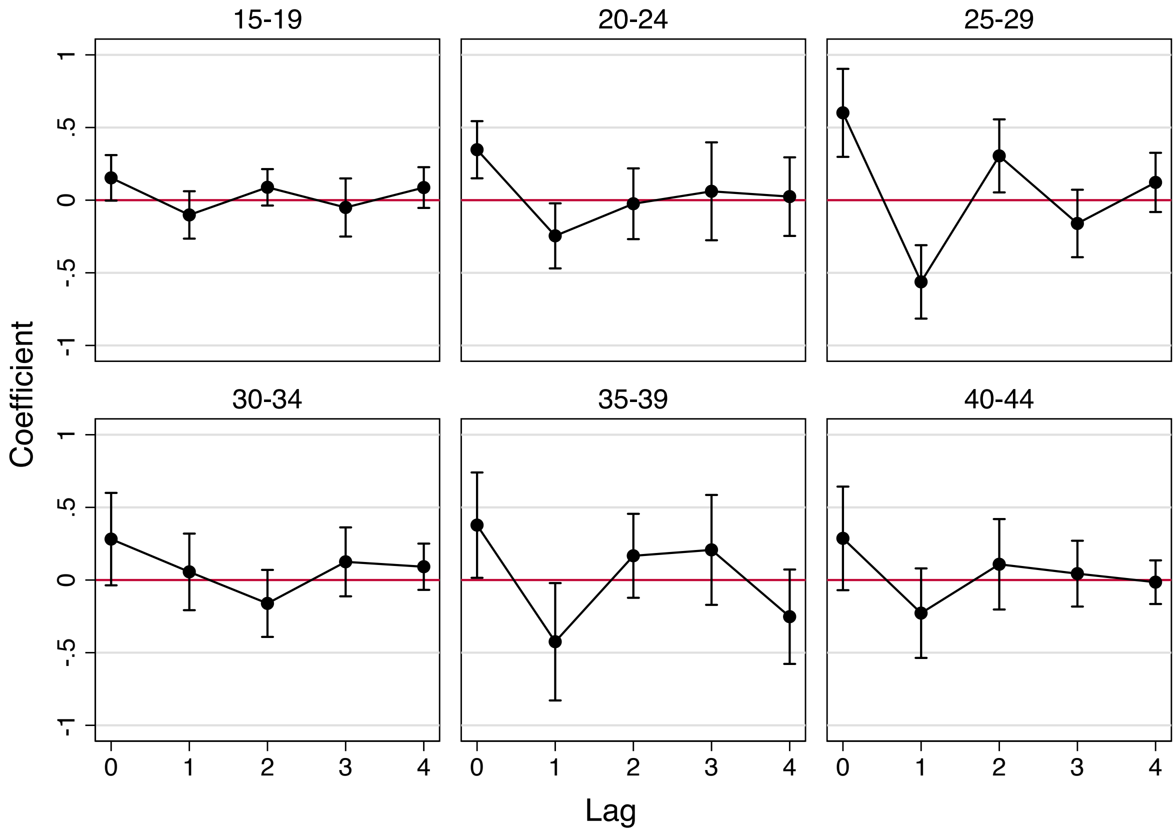
Notes: The figure compares results from different methods of computing the average annual rate of change. "20+ avg. ann. change" is the long-run model reported in Table 2, "20+ avg. obs. ann. change" uses the average of observed annual changes (leaving out gaps in the panel) instead of the slope of the annual trend. "Bartlett weights" downweights observations with longer recall periods. The remaining results use alternative minimum time horizons (15 and 25 years) to estimate the slope of the annual trend.

Figure A4: Non-Linear Estimates by Age Group, Short Run



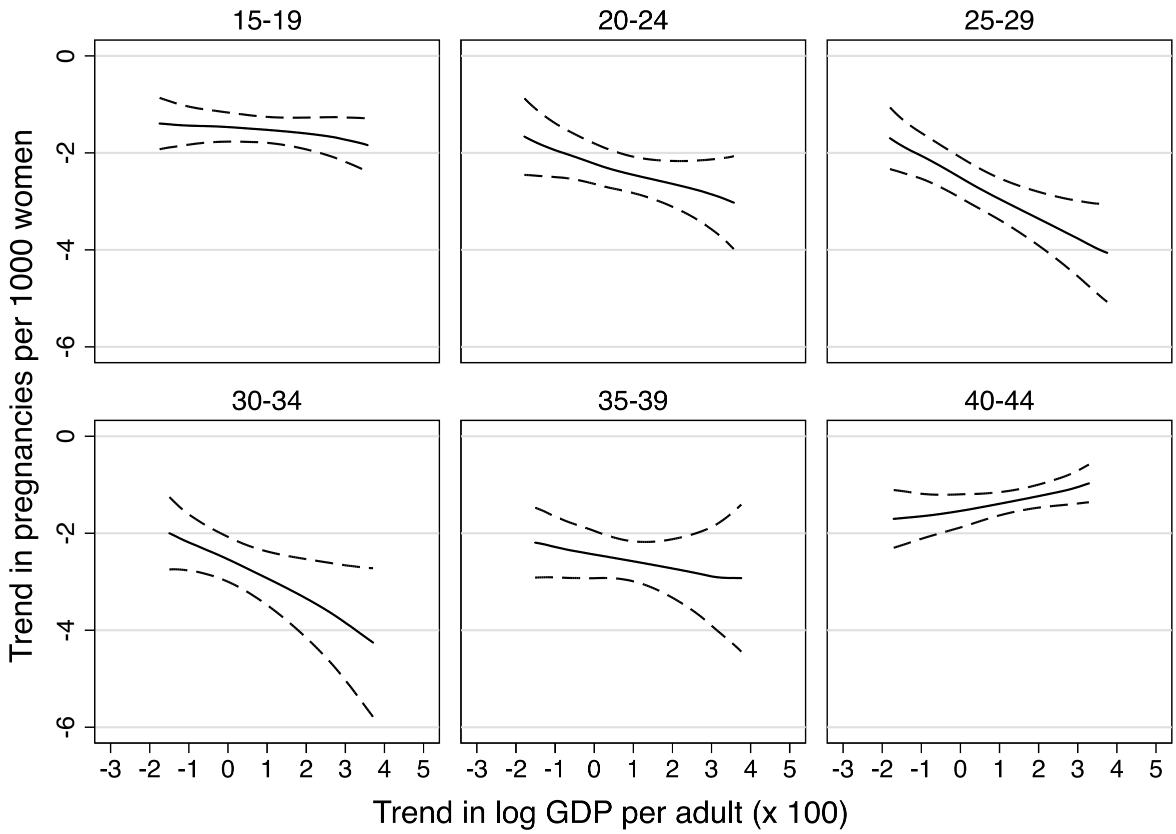
Note: Binned estimates reported in Figure 1, Panel A, here with 95% confidence intervals. Coefficients from regressions of annual changes in the age-specific conception rate on binned annual economic growth, controlling for country, year, and age fixed effects. Confidence intervals are clustered at the country level.

Figure A5: Distributed Lag Models by Age Group, Short Run



Note: Distributed lag model reported in Figure 3, Panel A, here with 95% confidence intervals. Coefficients from regressions of annual changes in the age-specific conception rate on current and lagged annual changes in 100 × log GDP per adult, controlling for country, year, and age fixed effects. Confidence intervals are clustered at the country level.

Figure A6: Non-Linear Estimates by Age Group, Long Run



Note: Local linear regression estimates with 95% confidence intervals. The dependent variable is the estimated trend in conception rates within a country-age cell, while the independent variable is the estimated trend in log GDP per adult in the same cell. Bandwidth equals 2, and confidence intervals are block bootstrapped at the country level.