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DIFFERENTIAL FERTILITY, HUMAN CAPITAL, AND DEVELOPMENT

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ABSTRACT

Using micro-data from 48 developing countries, I document a recent reversal in the income-fertility relationship and its aggregate implications. Before 1960, children from larger families had richer parents and obtained more education. By century's end, both patterns had reversed. Consequently, income differentials in fertility historically raised average education but now reduce it. While the reversal is unrelated to changes in GDP, women's work, sectoral composition, or health, half is attributable to rising aggregate education in the parents' generation. The results support a model in which rising skill returns lowered the minimum income at which parents invest in education.

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1 Introduction

Statisticians of the 19^{th} and early-20^{th} centuries expressed much concern about the negative correlations between fertility and a range of desirable attributes. Francis Galton, Karl Pearson, and Ronald Fisher, not to mention their many peers in the field of eugenics, all argued that the higher fertility rates of the poor implied the genetic deterioration of humankind (Kevles 1985). Over the next century, the pattern of 'differential fertility' between the rich and poor, the literate and illiterate, and the more and less intelligent caused alarm over the evolution of the distribution of traits.

Although eugenic arguments have gone out of vogue, social scientists have continued to study how differential fertility affects aggregate outcomes. Among modern economists, this interest dates back to Kuznets (1973), who suggested that differential fertility adversely affects both the distribution and the growth rate of income. A long line of research since then has formalized and further developed these theories.[1] At this literature's core is the observation that, in most present-day settings, wealthy parents have fewer children than poor parents, and they educate their children more. Compared to a population without heterogeneity in fertility rates, a population with greater fertility among the poor has a higher share of children from poor families, which lowers the average skill level. Some models also demonstrate how these fertility gaps can give rise to poverty traps, thus widening inequality.[2] Much of this work posits that the higher fertility of the poor can help explain the growth and income distribution experiences of developing countries over the 20^{th} century. This paper provides broad evidence that this claim cannot be true because, until quite recently, fertility increased with income throughout much of the developing world. Children from larger families therefore obtained *more* education.

At least since Becker (1960), economists have recognized that fertility may have once been positively correlated with income. But systematic evidence on the reversal of this cross-sectional relationship has emerged only recently, primarily for Western Europe several centuries ago (Hadeishi 2003; Clark 2007). What little evidence exists

[1] References include Althaus (1980), Dahan and Tsiddon (1998), Morand (1999), Galor and Moav (2002), Kremer and Chen (2002), de la Croix and Doepke (2003), Moav (2005), and Clark (2007).

[2] Empirically, Lam (1986) shows that the effect of differential fertility on inequality depends on the inequality metric. His finding does not overturn the general equilibrium reasoning of recent theories.

on currently poor countries is scattered, relying on small datasets from select parts of the world, mostly in Africa (Schultz 1986; Skirrbekk 2008). Similarly, despite bits of evidence that the relationship between sibship size and education may not have always been negative (Buchmann and Hannum 2001; Maralani 2008), this work has not identified generalizable patterns.

In the literature on the theory of fertility and its interaction with the macro-economy, the focus on the present regime with a negative income-fertility gradient is thus unsurprising. As its standard framework for the study of differential fertility, this literature uses a model that assigns a time cost to children and a goods cost to education: a setup that yields the negative gradient that is prevalent today. In Section 2 of the paper, I demonstrate that with the addition of a subsistence constraint or a goods cost of children, the same framework predicts that fertility increases with income among the poor. As such, in the early stages of development, children with more siblings come from richer families and obtain more education.

Drawing on extensive micro-data from 48 developing countries, I confirm these predictions by studying the evolution of two closely-related cross-sectional associations: (1) that between parental economic status (proxied by durable goods ownership) and fertility and (2) that between sibship size and education. In the not-too-distant past, richer families had more children, and children with more siblings obtained more education. Today, the opposite is true for both relationships. These findings have implications for theories of fertility and the demographic transition, as well as for understanding the role of differential fertility in the process of growth. In particular, until recently, differences in fertility decisions across families promoted the growth of the *per capita* human capital stock instead of slowing it.

To illustrate these findings, I construct two datasets from the Demographic and Health Surveys (DHS). For the first, I treat the survey respondents (who are women of childbearing age) as mothers, using fertility history data to construct two cross-sections of families from 20 countries in the 1986-1994 and 2006-2011 periods. In these data, respondents enumerate all of their children ever born, with information on survival status. Between the early and late periods, the relationship between parental durable goods ownership and the number of surviving children flipped from positive to negative in Africa and rural Asia; it was negative throughout in Latin America.

For the second dataset, I treat the DHS respondents as siblings, using sibling history data to retrospectively construct a longer panel of families from 42 countries. In these data, respondents report all children ever born to their mothers, again with information on survival status. Among earlier birth cohorts (mostly of the 1940s and 1950s), most countries show negative associations between the number of ever-born or surviving siblings and educational attainment. Among later birth cohorts (mostly of the 1980s), most countries show the opposite. The dates of the transition vary by setting, with Latin America roughly in the 1960s, Asia roughly in the 1970s, and Africa roughly in the 1980s. Taken together, the data suggest that in nearly all sample countries, both the income-fertility relationship and the sibsize-education relationship flipped from positive to negative. Indeed, although the DHS offers little data on childhood economic circumstance, three supplementary datasets (from Bangladesh, Indonesia, and Mexico) suggest that one can attribute much of the reversal in the sibsize-education relationship to the reversal of the income-fertility relationship.

I then quantify the changing effect of differential fertility on average educational attainment, relative to a thought experiment in which all families are forced to have the same family size. The theoretical framework shows that one can separate this effect into two components. The first reflects how the forced fertility policy would affect the composition of the population, while the second reflects how it would affect the distribution of education investment per child across families. I focus on the first component, which plays a larger role in theories of the aggregate effects of differential fertility, and which one can estimate by means of a simple reweighting procedure. The procedure compares actual average educational attainment with the (reweighted) average that would arise if all families had the same number of children, with no change to their educational attainment.

The results of the reweighting procedure are at odds with claims that differential fertility between rich and poor generally depresses average skill. Only in South Africa did differential fertility lower average education throughout the sample period. The remaining countries are split fairly evenly in two groups. In one, differential fertility elevated average education throughout the sample period, due to a consistently positive relationship between surviving sibship size and education. In the other, the influence of differential fertility changed over the sample period, typically starting positive and ending negative. The magnitudes are usually less than half a year of education: moderate in comparison to the nearly four-year increase in average educational attainment over the sample period. But they are meaningfully large relative to the level of average education in early cohorts. For women born during 1950-54, the reweighted average differs from the actual average by 15 percent.

Because these findings constitute the first systematic evidence of a reversal of differential fertility across the developing world, they are of interest independently of the mechanism mediating them. Nevertheless, to test alternative theories of the reversal, I assemble a country-by-birth cohort panel of sibsize-education coefficients. Net of country and cohort fixed effects, neither women's labor force participation, nor sectoral composition, nor GDP *per capita*, nor child mortality predicts the sibsize-education association. Rather, one variable can account for over half of the reversal of the sibsize-education association: the average educational attainment of the parents' generation.[3] These findings are broadly consistent with the theoretical framework in Section 2. However, because the reversal is uncorrelated with economic growth, its most likely cause is not a shift of the income distribution over the peak of a stable, hump-shaped income-fertility profile. Instead, a rising return to education may have lowered the income threshold at which families begin to invest in education, moving the peak of the income-fertility profile to the left. Indeed, although the fertility history data show a hump-shaped relationship between durable goods ownership and fertility in the late 1980s and early 1990s, the relationship is everywhere negatively sloped by the 2000s. The return-to-education theory is also consistent with the role of aggregate education; in many endogenous growth models (e.g., Becker et al. 1990; Galor and Weil 2000), aggregate human capital raises the individual return to education.

By documenting that patterns of differential fertility recently reversed, this paper makes contributions to several literatures. Most apparent is the connection with two empirical literatures: one on parental income and fertility, the other on sibship size and education. In both these literatures, existing evidence on positive income-fertility and sibsize-education associations is scattered, lacking a unifying framework.[4] This

[3]Male and female education have indistinguishable effects, suggesting that female empowerment does not explain the reversal.

[4]In a meta-analysis of 129 studies, Skirbekk (2008) finds a reversal in the income-fertility link. But the studies use varied methods on data from disparate settings, making the results difficult to interpret.

paper uncovers a common time path in which both associations flip from positive to negative. Building on a standard model of the growth literature, it provides a theoretical framework that explains the reversal and gives insight into its aggregate implications. Along these lines, the paper shows how cross-family heterogeneity in fertility historically increased average eduction but now largely decreases it. That finding adds to our understanding of how demography interacts with the macroeconomy and calls attention to how cross-sectional patterns can inform models of fertility decline. The basic time-series facts about fertility decline are overdetermined, so a more thorough treatment of changing heterogeneity within populations will help narrow the field of candidate theories of the demographic transition.

2 A Quality-Quantity Framework

This section demonstrates how a subsistence constraint or a goods cost of children influences the growth literature's standard theoretical framework for studying differential fertility. Given the paper's focus, I derive the model's main cross-sectional implications, rather than its intergenerational dynamics.

2.1 Setup

Parents maximize a utility function over their own consumption (c), the number of children (n), and human capital per child (h):

$$U(c,n,h) = \alpha \log(c) + (1-\alpha)\left(\log(n) + \beta \log(h)\right) \tag{1}$$

This utility function is standard in the literature on the demographic transition. $\alpha \in (0,1)$ indexes the weight the parents place on their own consumption relative to the combined quantity and quality of children, while $\beta \in (0,1)$ reflects the importance of quality relative to quantity. Child quality, or human capital, is determined by:

$$h(e) = \theta_0 + \theta_1 e \tag{2}$$

where e denotes education spending per child, and θ_0 and θ_1 are positive.[5] θ_0 is a human capital endowment (e.g., public school), while θ_1 is the return to education spending. One can view $h(e)$ as a child's earnings in adulthood or as some broader measure of human capital. Irrespective of human capital, each child costs $\tau \in (0,1)$ units of time and $\kappa \geq 0$ goods. These costs represent the minimum activities (e.g., pregnancy, child care) and goods (e.g., food, clothing) required for each child.

Parents work at wage w with a time endowment of 1, so the budget constraint is:

$$c + \kappa n + ne \leq w(1 - \tau n) \tag{3}$$

They also face a subsistence constraint: c must exceed $\tilde{c} \geq 0$. The framework allows the child goods cost and the subsistence level to be zero, in which case it reduces to the models of De la Croix and Doepke (2003) and Moav (2005). It focuses on wages rather than income, but qualitative predictions for the two variables are the same.

2.2 Optimal Fertility and Education Spending

This setup yields closed-form solutions for optimal fertility and education spending. If the wage is below the threshold $\tilde{w} = \frac{1}{\tau}\left(\frac{\theta_0/\theta_1}{\beta} - \kappa\right)$, parents are content with the human capital endowment θ_0 and choose a corner solution with no education spending. If $w \geq \tilde{w}$, education spending per child rises linearly in the wage:

$$e_w^* = \begin{cases} 0 & \text{if } w < \tilde{w} \\ \frac{\beta(\kappa+\tau w)-\theta_0/\theta_1}{1-\beta} & \text{if } w \geq \tilde{w} \end{cases} \tag{4}$$

The subsistence level $\tilde{c}$ plays no role in determining education spending per child.

The subsistence constraint does influence fertility choice, however:

$$n_w^* = \begin{cases} \frac{w-\tilde{c}}{\kappa+\tau w} & \text{if } w < \min\left(\frac{\tilde{c}}{\alpha}, \tilde{w}\right) \\ \frac{(1-\alpha)w}{\kappa+\tau w} & \text{if } \frac{\tilde{c}}{\alpha} \leq w < \tilde{w} \\ \frac{(1-\beta)(w-\tilde{c})}{\kappa-\theta_0/\theta_1+\tau w} & \text{if } \tilde{w} \leq w < \frac{\tilde{c}}{\alpha} \\ \frac{(1-\alpha)(1-\beta)w}{\kappa-\theta_0/\theta_1+\tau w} & \text{if } w \geq \max\left(\frac{\tilde{c}}{\alpha}, \tilde{w}\right) \end{cases} \tag{5}$$

[5] To add concavity, one can set $h(e) = (\theta_0 + \theta_1 e)^{\sigma}$, $\sigma \in (0,1)$, without affecting the results.

The first line of Equation (5) corresponds to the case in which the parents are both subsistence constrained and at an education corner solution. After consuming $\tilde{c}$, they spend all of their remaining full income w on child quantity, so fertility increases with the wage. The next two lines deal with the cases in which $\frac{\tilde{c}}{\alpha} < \tilde{w}$ and $\tilde{w} < \frac{\tilde{c}}{\alpha}$, respectively. In the second line, the subsistence constraint no longer binds, but the parents remain at an education corner solution. They devote αw to their own consumption and the remainder to child quantity, so fertility is increasing in the wage if $\kappa > 0$ and constant if $\kappa = 0$. In the third line, the subsistence constraint binds, but the parents now choose an education interior solution, making the comparative static ambiguous: $\frac{dn_w^*}{dw} \gtreqless 0$ if and only if $\kappa \gtreqless \frac{\theta_0}{\theta_1} - \tau\tilde{c}$. It is also ambiguous in the final line, in which the parents are constrained by neither the subsistence constraint nor the lower bound on education spending: $\frac{dn_w^*}{dw} \gtreqless 0$ if and only if $\kappa \gtreqless \frac{\theta_0}{\theta_1}$.

To summarize, either a subsistence constraint or a goods cost of children guarantees a hump-shaped relationship between wages and fertility, so long as the goods cost is not too large.[6] At low wage levels, fertility increases with the wage if $\kappa > 0$ or $\tilde{c} > 0$; at high wage levels, it decreases in the wage if $\kappa < \theta_0/\theta_1$. For illustration, Figure 1 graphs n_w^* and e_w^* against w for the case in which $0 < \frac{\tilde{c}}{\alpha} < \tilde{w}$ and $0 < \kappa < \theta_0/\theta_1$. The n_w^* profile follows a hump-shape, with fertility low among the poor, high among parents with intermediate wages, and somewhat lower among the rich, converging toward the asymptote $(1-\alpha)(1-\beta)/\tau$. These results suggest that the wage-fertility association is positive in the early stages of economic development, but that broad-based gains in living standards tend to make it negative.

The parameters of the model may also change over the course of development, with implications for the shape of the wage-fertility relationship. Two changes are relevant for the empirical work below. First, as Malthus (1826) once argued, the rise of anti-poverty programs may alleviate the subsistence constraint, thus decreasing the amount $\tilde{c}$ that parents need to devote to their own consumption. Second, as emphasized by Unified Growth Theory (Galor 2011), the return to education spending θ_1 may increase, thus decreasing $\tilde{w}$. These forces shift the maximum of the n_w^* profile to the left, so they also tend to make the wage-fertility association negative.

[6] Absent a subsistence constraint or a goods cost of children, fertility is invariant to the wage when $w < \tilde{w}$ and decesasing in the wage when $w \geq \tilde{w}$. Because children bear only a time cost and education bears only a goods cost, the price of child quantity is relatively high for higher-wage families.

2.3 Implications for Average Human Capital

To characterize the effect of differential fertility on average human capital, assume a wage distribution $F(w)$ on $[\underline{w},\overline{w}]$, and consider a policy forcing all couples to have $\tilde{n}$ children.[7] The effect of differential fertility is the difference between average human capital under free fertility and average human capital under forced fertility. Under forced fertility level $\tilde{n}$, parents with wage w choose education spending as follows:

$$e_w^{\tilde{n}} = \begin{cases} \frac{w-\tilde{c}}{\tilde{n}} - \kappa - \tau w & \text{if } w < \frac{\tilde{c}/\alpha+\kappa\tilde{n}}{1-\tau\tilde{n}} \\ \frac{(\beta-\alpha\beta)\left(\frac{w}{\tilde{n}}-\kappa-\tau w\right)-\alpha\theta_0/\theta_1}{\alpha+\beta-\alpha\beta} & \text{if } w \geq \frac{\tilde{c}/\alpha+\kappa\tilde{n}}{1-\tau\tilde{n}} \end{cases} \tag{6}$$

Consistent with a quality-quantity tradeoff, education spending decreases in $\tilde{n}$. Note that $e_w^{n_w^*}$ equals e_w^*, optimal education spending under free fertility.

For wage distribution F and forced fertility level $\tilde{n}$, the total effect of differential fertility on average human capital is thus:

$$\Delta_{tot}(F,\tilde{n}) = \frac{\int h(e_w^*)\, n_w^* dF(w)}{\int n_w^* dF(w)} - \frac{\int h\left(e_w^{\tilde{n}}\right) \tilde{n} dF(w)}{\tilde{n}} \tag{7}$$

On the right-hand side of the equation, the first and second expressions equal average human capital under free and forced fertility, respectively. To average across children rather than families, both expressions reweight the wage distribution by the factor $\frac{n}{E[n]}$. In the second expression, all families have the same fertility level, so this factor equals 1. Although $\Delta_{tot}(F,\tilde{n})$ is relevant to interventions like China's one child policy, coercive fertility polices are rare, so it has few real-world applications.

One can decompose $\Delta_{tot}(F,\tilde{n})$ into two quantities, one of which does not depend on a counterfactual policy. To obtain this decomposition, add and subtract $\int h(e_w^*)\, dF(w)$, average human capital *across families*, to the right-hand side of Equation (7):

$$\Delta_{tot}(F,\tilde{n}) = \underbrace{\int \left(\frac{n_w^*}{\int n_\upsilon^* dF(\upsilon)} - 1\right) h(e_w^*)\, dF(w)}_{\Delta_{comp}(F)} + \underbrace{\int \left\{h(e_w^*) - h\left(e_w^{\tilde{n}}\right)\right\} dF(w)}_{\Delta_{adj}(F,\tilde{n})} \tag{8}$$

[7] For simplicity, assume $\tilde{n} < \frac{w-\tilde{c}}{w\tau-\kappa}$, so that the forced level of fertility is not so high that it prevents parents from meeting the subsistence constraint.

where υ is a dummy of integration. $\Delta_{comp}(F)$ is the *composition effect* of differential fertility, measuring how average human capital *across children* differs between the free fertility optimum and the counterfactual in which all families have an equal number of children but maintain the per child educational investments that were optimal under free fertility. Because this counterfactual involves no re-optimization, the composition effect is invariant to $\tilde{n}$. $\Delta_{adj}(F,\tilde{n})$ is the *adjustment effect* of differential fertility, measuring how average human capital *across families* changes in response to a policy shift from free fertility to forced fertility level $\tilde{n}$. This component depends crucially on $\tilde{n}$. Under a policy forcing the lowest observed fertility rate on all parents, the adjustment effect would be positive; if the policy instead forced the highest observed fertility rate, the adjustment effect would be negative.

The empirical work focuses on the composition effect because it solely reflects the joint distribution of quantity and quality investments, rather than arbitrarily-defined counterfactual policies. Assuming a positive subsistence level and a small goods cost of children, several properties of the composition effect are apparent. If $\overline{w} < \tilde{w}$, so that all parents make no educational investments, then $\Delta_{comp}(F) = 0$. As $\overline{w}$ rises above $\tilde{w}$, $\Delta_{comp}(F)$ turns positive because fertility rates are highest in the small share of parents with positive education spending. Further rightward shifts in F (or leftward shifts in $\tilde{w}$) lead to more mass in the domain in which $\frac{dn_w^*}{dw} < 0$, eventually turning $\Delta_{comp}(F)$ negative. Indeed, if $\underline{w} > \max\left(\frac{\tilde{c}}{\alpha}, \tilde{w}\right)$, so that fertility decreases in the wage across the entire support of F, then $\Delta_{comp}(F)$ is unambiguously negative. These results suggest that in the early stages of economic development—when most are subsistence constrained or at an education spending corner solution, but the wealthy few educate their children—the composition effect is positive. But with broad-based gains in living standards, increases in the return to education spending, or expansions in anti-poverty programs, the composition effect turns negative.

3 Data on Two Generations of Sibships

Using data from the Demographic and Health Surveys (DHS), I construct two generations of sibships by viewing respondents as both mothers and daughters. Carried out in over 90 countries over the past three decades, these surveys interview nationallyrepresentative samples of women of childbearing age (generally 15-49). All include questions about the respondent's educational attainment and children; some also ask about household durable goods ownership or the respondent's siblings.

3.1 DHS Fertility Histories

The first set of analyses draws on the fertility histories, in which respondents list all of their children ever born, with information on survival. To avoid the complicated task of disentangling cohort effects from changes in the timing of childbearing, I focus on women at least 45 years old and interpret their numbers of children as completed fertility. The focus on older women also has the advantage of capturing cohorts of mothers more likely to be in the early regime in which fertility is increasing in income. I compare results from two time periods, pre-1995 and post-2005, and only include countries with survey data from both periods, leaving me with 62,146 women from 46 surveys in 20 countries.

The DHS does not collect data on wages or income, so a direct test of the theoretical framework is not possible. However, the surveys do include questions on household ownership of several durable goods, which are useful from two perspectives. First, durables ownership can serve as a proxy for broader measures of consumption or income. In the presence of transitory shocks, consumption and income are not *a priori* superior to durable goods indices as measures of long-term income, which is probably the more relevant determinant of childbearing decisions. Alternatively, one can more explicitly view durable goods ownership as an outcome of optimization. If one added $\log(d)$ to the utility function in the theoretical framework, with d representing durables, then optimal durables ownership would bear a hump-shaped relationship with fertility and a weakly increasing relationship with education.[8]

Existing work on the DHS has drawn extensively on durable goods ownership to measure economic status, much of it using the method proposed by Filmer and Pritchett (2001), which takes the first principal component of a vector of variables measuring housing conditions and ownership of several durable goods. I modify this approach in two ways. First, I only use data on ownership of five durable goods:

[8] In his original work on fertility, Becker (1960) drew a close link between the demand for children and the demand for (other) durable goods.

radio, television, refrigerator, motorcycle, and car.[9] By not incorporating measures of housing conditions, I avoid the tasks of determining whether certain conditions (e.g., access to piped water) are communally determined and whether they directly influence fertility. Second, rather than using principal components analysis, I take the sum of the ownership indicators. The resulting index of durables ownership is both transparent and comparable across time and space, notwithstanding concerns about changes in relative prices. One DHS, the 1994 survey from Indonesia, included an expenditure module; in that survey, a regression of the durables index on log per capita expenditure yields a coefficient of 1.01 (S.E. = 0.04).

3.2 DHS Sibling Histories

The DHS began administering a sibling history module in the late 1980s for the purpose of estimating maternal mortality rates in settings with poor or absent vital registration systems. The module asks respondents to list all children ever born to their biological mothers, with information on sex, year of birth, and year of death if no longer alive. Analyses of maternal mortality and all-cause adult mortality have since then drawn extensively on DHS data (e.g., Obermeyer et al. 2010). However, the sibling history data also offer a window into the sibling structure that adult women experienced as children.

As of December 2012, data from 89 DHS's with full sibling histories were in the public domain. Of these, seven (from Bangladesh, Indonesia, Jordan, and Nepal) included only ever-married women, introducing concerns about selection bias. From these surveys, I only include age groups in which the rate of ever marriage is at least 95 percent. Therefore, I include women over 30 from the relevant surveys in Bangladesh and Nepal, but I discard the 5 surveys from Indonesia and Jordan, where female marriage rates are lower. Nepal has two surveys with sibling histories, one of ever-married women in 1996 and one of all women in 2006. I restrict the 1996 sample to women over 30, but I include all respondents to the 2006 survey. I also discard data from the 1989 Bolivia DHS and the 1999 Nigeria DHS due to irregularities in the sibling history data, leaving 82 surveys for analysis. Africa is overrepresented, a

[9] Many surveys ask about bicycle ownership, but I omit it because it may be endogenous to the presence of children.

consequence of the near absence of systematic data on adult mortality in the continent prior to the entrance of the DHS. To exclude respondents who have not finished schooling or whose mothers have not completed childbearing, I drop data on women less than 20 years old, leaving 803,527 women born between 1942 and 1989.

3.3 Supplementary Surveys

The DHS data are useful in their breadth but suffer from two major shortcomings. The most obvious is their omission of men, for whom the relationship of interest may be different. Additionally, they offer little information on aspects of the respondent's childhood environment, such as the income or education of her parents. To supplement the DHS on these two fronts, I draw on three supplementary surveys in the Appendix: the Indonesia Family Life Survey, the Matlab Health and Socioeconomic Survey, and the Mexico Family Life Survey. All three surveys include questions about surviving siblings and parental characteristics.

4 Changing Cross-Sectional Fertility Patterns

This section documents the evolution of differential fertility in developing countries over the second half of the twentieth century. In all of the analyses, I first separate the sample into country-by-period cells and then estimate a mean or regression coefficient within each cell.[10] For any cross-country results, I then perform unweighted analyses of the cell-level statistics.

4.1 Durable Goods Ownership and Fertility

To assess the changing association of durable goods ownership and fertility, I use fertility history data to estimate separate country-level regressions for survey respondents aged 45-49 in the early (1986-1994) and late (2006-2011) DHS periods. For woman i in county c and period t, I run:

$$fertility_{ict} = \delta_{ct} + \gamma_{ct} index_{ict} + X'_{ict}\lambda_{ct} + \varepsilon_{ict} \tag{9}$$

[10]The analyses use sampling weights, but the results are similar without them.

where $fertility_{ict}$ denotes the woman's number of children (ever born or surviving), $index_{ict}$ denotes the durable goods ownership index (which varies between 0 and 5), and the vector X_{ict} contains age indicators and survey year indicators.

The main results for both ever-born fertility and surviving fertility appear in Table 1, which shows averages of the country-specific coefficients at the continent level. Panel A pools urban and rural areas, showing results both with and without controlling for an urban residence indicator. Panels B and C report results for solely urban and solely rural areas, respectively. A cross signifies that the late-period coefficient differs significantly from the early-period average coefficient. To aid in the interpretation of the results, Appendix Table 1 shows continent-by-period means and standard deviations of the relevant variables.

Table 1 reveals a reversal in the relationship of durable goods ownership and surviving fertility: certainly for Africa and to some extent for Asia, but not for Latin America.[11] In Africa, controlling for urban residence (Panel A2), each additional durable good is associated with one-fifth *more* surviving children in the early period but one-fifth *fewer* children in the late period. This flip is especially pronounced in rural areas (Panel C). Indeed, the same patterns hold in rural areas of the Asian countries in the sample, although not in urban areas of these countries. In the full Asian sample, controlling for urban residence, the durables index is uncorrelated with surviving fertility during the early period, but the association turns negative by the late period. All of these inter-period changes in coefficients are statistically significant at the 5 percent level. The same patterns do not generally hold in Latin America, where the durable goods index negatively predicts surviving fertility in both the early and late periods. Nevertheless, in rural areas within Latin America, the relationship becomes significantly more negative over time. These results may suggest a shared process that operates at different times across and within countries: visiting urban areas before rural, and visiting Latin America before Asia and Africa.

When one counts all children ever born instead of only those that survived, the results change in predictable ways. Survival rates are positively related to parental income and education throughout the sample period, which makes the ever-born coefficients more negative than the surviving coefficients. Indeed, the durables index is

[11] Appendix Table 2 shows that these results are robust to the inclusion of paternal and maternal education as covariates.

negatively correlated with ever-born fertility in all regions and time periods, although the relationship is small and statistically insignificant for rural Africa in the early period. Throughout Africa and Asia, the relationship becomes more negative between the early and late periods. Again, the sibling history results will shed light on whether ever-born fertility was positively associated with income in an earlier time.

Non-parametric estimation can shed light on whether the linear regression coefficients mask the hump shape predicted by the theory. For a first look at this issue using varied measures of household consumption, Figure 2 takes advantage of Indonesia's 1994 survey, which included both durable goods questions and an expenditure module. Consistent with the theoretical framework, durable goods ownership, log total household expenditures, and log household expenditures per adult all display hump-shaped relationships with both ever-born and surviving fertility.

Figure 3 replicates the durable goods graph for the full fertility history sample. The plots show clear evidence of a hump shape in the 1986-94 period, especially for counts of surviving children. However, outside Africa, the hump dissipates by the 2006-11 period. This finding is inconsistent with a stable, hump-shaped relationship between income and fertility. Relative to many other goods, the prices of the durable goods included in the index most likely *decreased* over the relevant period, so that durable goods ownership diffused down the income distribution. In that case, one would expect the peak of the relationship to shift to the right rather than the left. However, as discussed in the theoretical framework, expansions in public assistance and increases in the return to schooling move the peak to the left. These forces may explain the disappearance of the hump.

4.2 Sibship Size and Educational Attainment

The fertility history results provide evidence of a reversal in the relationship between durable goods ownership and surviving fertility in Africa and rural Asia, but they leave several questions unanswered. Did the same reversal occur for counts of ever-born children at some earlier date? Did it occur in Latin America? The sibling histories offer a window onto the answers to these questions for birth cohorts going back to the 1940s. Unfortunately, the DHS collects very little data on economic conditions in childhood. However, the theoretical framework suggests that we can infer the evolution of the income-fertility association from changes in the relationship between sibship size and education. The sibsize-education link is also directly relevant for assessing the effect of differential fertility on the skill distribution. I estimate separate country-level regressions for women in 5-year birth cohorts from 1945-1949 to 1985-1989.[12] For woman i born in county c and time period t, I run:

$$highest\,grade_{ict} = \delta_{ct} + \gamma_{ct} sibsize_{ict} + \varepsilon_{ict} \tag{10}$$

where $highest\,grade_{ict}$ denotes her schooling and $sibsize_{ict}$ denotes her sibship size.[13]

Figure 4, which displays estimates of γ_{ct} over time within each country, makes clear that attempts to characterize the sibsize-education relationship as generally negative miss a pervasive feature of recent demographic history. Both the ever-born sibling and the surviving sibling coefficients tend to decrease across successive birth cohorts. For earlier birth cohorts, most coefficients are significantly positive, while for the latest birth cohorts, few coefficients are significantly positive, and many are significantly negative. Consistent with the fertility history results, this reversal in the sibsize-education relationship occurs earliest in Latin America, followed soon thereafter by several countries in Asia. In Africa, the reversal has been quite recent, and several countries remain in the pre-reversal regime. To put these estimates in context, Appendix Figure 1 plots trends in average education, showing several-year increases in average education in most countries. In absolute value, γ_{ct} is small when average education is low.

These results leave two issues unaddressed: birth order effects and gender heterogeneity. Birth order is a concern because children of high birth orders necessarily come from large families. Given evidence that birth order affects educational attainment (Steelman et al. 2002; Black et al. 2005), researchers often control for birth order in estimating the effect of family size on educational attainment. However, the present paper is concerned not with causal effects but with equilibrium differences between large and small families, making regression adjustment unnecessary. Birth order effects are but one reason for the different outcomes of children from large and small families. On gender heterogeneity, although the DHS only gathers sibling

[12] For precision, I omit cells with fewer than 200 observations, representing 2.5 percent of all cells.
[13] The results are robust to the inclusion of birth year fixed effects.

history data from women, the supplementary surveys from Bangladesh, Indonesia, and Mexico interview both genders. In Appendix Table 3, all three supplementary surveys show declining sibsize-education relationships for both genders.

4.3 Connecting the Results

The fertility history results seem to contain the last phases of the global transition to a negative relationship between income and fertility, while the sibling history results point to a widespread shift of the sibsize-education link from positive to negative. While the two phenomena seem connected, the absence of childhood background characteristics in the DHS prevents examination of this issue. The supplementary surveys include data on paternal education, which can shed some light on the role of economic resources in childhood. Using these data, Appendix Table 4 compares sibsize-education coefficients from regressions that do and do not control for paternal education.[14] In the specification that controls for paternal education, decreases in the coefficients across successive birth cohorts are muted by at least one half. The evolution of the sibsize-education relationship has much to do with a changing relationship between paternal education and sibship size.

5 Differential Fertility and Average Human Capital

The results so far suggest that differential fertility once promoted human capital accumulation rather than hindering it. This section estimates the changing composition effect of differential fertility on average education.

Recall from the theory section that the composition effect is:

$$\Delta_{comp}(F) = \int \left(\frac{n_w^*}{\int n_\upsilon^* dF(\upsilon)} - 1 \right) h(e_w^*)\, dF(w) \tag{11}$$

This expression integrates over the parental wage distribution, but I only observe siblings, with little information about their parents. Applying the law of iterated expectations, I thus rewrite the composition effect over the distribution of surviving

[14]The results are similar in unreported analyses that also include maternal education as a covariate.

sibship sizes:

$$\Delta_{comp}(F) = \sum_{k=1}^{K} \left(\eta_k - \frac{\eta_k/k}{\sum_{l=1}^{K} \eta_l/l} \right) \mu_k \tag{12}$$

where K is the maximum possible sibship size, η_k is the share of the individuals from surviving sibships of size k, and μ_k is the mean human capital of individuals from sibships of size k. Inside the parentheses, the term η_k weights the sample to give mean human capital across individuals, while the term $\frac{\eta_k/k}{\sum_{l=1}^{K} \eta_l/l}$ reweights the sample to give mean human capital across families. Importantly, this expression captures *any* composition effect of heterogeneity in fertility and skill investment, not just the income heterogeneity specific to the model in Section 2.

I use the empirical analogues of α_k and μ_k to estimate $\hat{\Delta}_{comp}$ and then use the delta method to estimate the variance of $\hat{\Delta}_{comp}$. In the interest of simplicity, I measure human capital as highest grade completed. For successive 5-year birth cohorts within each country, Figure 5 displays estimates of the composition effect of differential fertility on average educational attainment.[15]

The results overturn the conventional wisdom that variation in fertility over the income distribution tends to lower average education. In some countries, predominantly African, differential fertility increased average educational attainment throughout the sample period. These countries have not transitioned to the regime in which surviving sibship size and education are negatively correlated. Opposite these countries is South Africa, where the effect of differential fertility was negative throughout almost the entire sample period. The remaining countries have undergone a transition from a regime in which differential fertility promoted the growth of human capital to a regime in which differential fertility depressed it. For two compelling examples, consider the Andean nations of Bolivia and Peru. For the 1945-9 cohort, differential fertility increased average education by 0.3 to 0.5 years in both countries. In contrast, for the 1985-9 cohort, differential fertility reduced average education by 0.5 years.

Are these magnitudes large or small? The answer depends on whether one evaluates them relative to the *increase* in education over the sample period or relative to the historical *level* of education. On average, the 1985-9 cohorts have 3.7 more years of

[15] To examine whether a single education level drives the results in Figure 5, Appendix Figure 2 estimates composition effects on shares of each cohort with 0, 1-5, 6-8, and 9+ years of education. The shifting composition effects are visible at all levels, from 0 years through 9+ years.

education than the 1945-9 cohorts.[16] The largest estimated composition effects are ± 0.6, and the average within-country change in these effects between 1945-9 and 1985-9 is -0.17. Therefore, the shift from a positive to a negative sibsize-education relationship did not have a large effect on the evolution of average educational attainment. But relative to the level of average educational attainment, the composition effect is reasonably large for early cohorts. For the 1950-4 cohort, the composition effect was on average 15 percent of mean education. As mean education rose, the relative magnitude of the differential fertility effect shrank: for the 1985-9 cohort, the effect of differential fertility on mean education was on average 4 percent of the cohort's mean education.

6 Explaining the Reversal

The reversal of differential fertility in the developing world occurred during a half-century that included much economic and demographic change. Although Section 2 suggests a compelling theory for the change, the existing literature suggests some alternatives. This section lists forces often associated with the demographic transition and explores their possible roles in the reversal. Although the discussion is theoretical, it aims to make predictions regarding the aggregate determinants of γ_{ct}.

Income Growth Section 2's theoretical framework suggests that broad-based income gains pushed families over the hump of the non-monotonic relationship between income and fertility. Under this hypothesis, the reversal of γ_{ct} should be associated with rising GDP *per capita* and with rising average adult education. The relative predictive power of these two variables depends on their association with incomes throughout the income distribution. The hypothesis also predicts that the reversal of γ_{ct} will be associated with rising average educational investment.

Human Capital The rise in the demand for and supply of schooling plays a key role in many models of the transition from Malthusian stagnation to growth. In Sec-

[16] This claim is based on a regression of cohort average education on country and cohort indicators. The coefficient on the 1985-9 cohort indicator is 3.7, indicating that the 1985-9 cohorts have 3.7 more years of education than the omitted category, the 1945-9 cohorts.

tion 2's theoretical framework, an increase in the return to education spending θ_1 can move the peak of the wage-fertility profile to the left, which is consistent with the reversal and with the disappearance of the hump in Figure 3. Unlike the income growth hypothesis, the return-to-schooling hypothesis does not predict a role for GDP *per capita*. However, if increases in aggregate human capital push up the return to schooling, as in the models of Becker et al. (1990) and Galor and Weil (2000), then the hypothesis predicts the decline of γ_{ct} to be associated with rising average adult education. It will also be associated with rising average educational investment and declining average family size.

The human capital endowment θ_0 can also factor into the explanation; an increase in that parameter, perhaps from an expansion in compulsory public schooling, makes the wage-fertility relationship more negative. But optimal fertility weakly increases in θ_0. As a result, we can reject one of the two human capital explanations based on whether the reversal of γ_{ct} is associated with rising or falling average family size.

Children's Work A related issue is the falling prevalence of child labor, which in the theoretical framework has similar consequences to a rising return to education spending. Some of the decline in child labor may actually be the result of increases in skill returns. Some might also be due to new sanctions against child labor, which one could characterize as an increase in the goods cost of children κ. Just as with an increase in the return to education spending, an increase in the goods cost of children decreases the wage threshold at which families start to spend on education, which can shift the peak of the wage-fertility profile to the left. This mechanism is complementary to the return-to-schooling hypothesis.[17]

Women's Work At least as likely an explanation as children's work is women's work. The reasoning is similar to that of Galor and Weil (1996), who argue that skill-biased technological progress increased women's labor productivity over the long run, eventually inducing greater women's labor force participation and lowering

[17] In another version of the child labor theory, family labor is cheaper than outside labor, so that landed agricultural households have increased demand for children as laborers. If landed agricultural households are drawn from the center of the income distribution, their demand for child labor can generate a hump-shaped income-fertility relationship.

fertility due to the increased opportunity cost of childbearing. They consider neither quality investments nor cross-sectional heterogeneity, but such extensions are natural. In Section 2's framework, one cannot generate a negative wage-fertility association without assuming a positive opportunity cost of childcare time.

This explanation runs up against the empirical reality, originally documented by Goldin (1995), that women's labor force participation follows a u-shape over the course of economic development.[18] Rates of women's labor force participation were high in Africa throughout the sample period, despite a positive relationship between income and fertility. But a closer reading of Goldin (1995) suggests that in the early stages of development, when labor is mostly agricultural, women's work is compatible with child rearing. Women's labor force participation then decreases when manufacturing predominates and increases with the emergence of the service sector. Unlike agricultural work, service jobs compete with childbearing. If women's opportunity cost of time explains the reversal, then the emergence of the service sector must also play a key role.

Child Mortality The decline of child mortality is also central to many theories of fertility decline, but it is unlikely to explain the change in fertility regimes observed in this paper. Because the bulk of mortality decline has occurred for children younger than school-starting age, one can think of a it as a reduction in the quantity costs of surviving children. In Section 2, a decline in the goods costs of children can make the slope of the wage-fertility relationship more negative at high wages, although it also moves the peak of the relationship to the right. In any case, as in the Barro-Becker model (1989), reductions either the costs of child quantity lead to higher optimal fertility and lower optimal schooling investment, which appears counterfactual. If child mortality is behind the reversal, then the theoretical framework predicts that the decline of γ_{ct} will be associated with rising average family size, declining education spending, and declining child mortality.

Preference Change In interpreting the changing cross-sectional patterns, many non-economists would think first of preferences. Several theories fertility decline

[18] Also see Mammen and Paxson (1998), and Olivetti (2012).

(Caldwell 1980, 1982; Casterline 2001) posit changes in beliefs and norms regarding child-rearing. Some versions of these theories could explain the observed regime change. Consider the introduction of new 'Western' norms that increase the relative importance of child quality in the utility function (β), raising optimal education and lowering optimal fertility. If these new norms affect the richest (or most educated) families most strongly, then the income-fertility relationship could flip from positive to negative, starting at the right tail of the income distribution. Caldwell (1980, 1982) assigns much importance to mass education in altering childbearing norms, thus predicting a relationship between γ_{ct} and average adult education. However, without further structure, the theory is otherwise difficult to test.

A more testable version associates the diffusion of new norms with the empowerment of women (Duflo 2012). If women have lower β's than men, and if women of higher income or education make the earliest gains in household bargaining power, then richer households will be the first to transition to low fertility. This reasoning predicts that fermale empowerment measures will be negatively associated with γ_{ct}.

Other versions of preference-change theory are also consistent with the data. Galor and Moav (2002, see also Clark 2007) develop one such version by combining β-heterogeneity with a subsistence constraint. Their model's evolutionary dynamics generate a reversal of the wage-fertility elasticity from positive to negative, just as observed in this paper. Family dynasties with high β's accumulate more human capital and therefore become richer than their low-β counterparts. Early in the process of development, the subsistence constraint binds for the poorer, low-β types, so that the high-β types choose higher fertility in addition to higher investment per child. This differential fertility pushes up the average skill level in the population, generating technological progress that gradually pushes low-β families over the subsistence constraint. At that point, the poorer, low-β types transition to high fertility, leading to a negative wage-fertility relationship. Here too, rising aggregate human capital will be associated with a reversal in γ_{ct}. Because Galor and Moav's model depends crucially on a subsistence constraint, it is complementary to Section 2's framework.

Intergenerational Wealth Transfers A separate class of theories, which does not fit into the framework above, emphasizes upward intergenerational transfers from children to parents, in the form of child labor or old-age support.[19] Caldwell (1982) emphasizes how the expansion of schools alters child-rearing norms, so that parents come to view children as net recipients of, rather than net contributors to, household resources. This model bears similarities with other theories of changing preferences. Following a different thread in Caldwell's work, Boldrin and Jones (2002) study parental behavior when old-age security is the primary motive for childbearing. In their framework, financial deepening could flip the income-fertility relationship if wealthy families substituted other savings vehicles for children. But this reasoning gives no account for why the decreases in quantity investment would be accompanied by increases in quality investment. Additionally, as stressed by Galor (2011), wealthier couples typically have access to a wider variety of savings vehicles before the fertility transition. Finally, Lee (2000) argues that data from no society suggest a net upward flow of resources across generations, unless one counts the pension systems of rich countries.

Contraception Advocates of family planning might instead emphasize the uneven adoption of effective contraceptive technology (Potts 1997). From this perspective, the currently negative relationship between income and fertility is due to an unmet need for contraception among the poor. But a theory of this type fails to account for the early regime during which fertility increases in income. One possibility is that women from richer households have a higher biological capacity to bear children due to their better health. In this case, broad-based health improvements would decrease the relationship between income and fertility.

7 Aggregate Determinants of the Reversal

With an eye to the explanations described in Sections 2 and 6, this section estimates how several economic and demographic aggregates relate to the sibsize-education link. I focus on the sibsize-education link rather than the income-fertility link because the former offers a longer time horizon and is more precisely estimated at the country level. Additionally, I only show results for the surviving sibship size coefficients

[19] See Cain (1983), Nugent (1985), Ehrlich and Lui (1991), and Morand (1999).

because they bear a closer link to theoretical framework and because they are directly relevant to the composition effect. Unreported results for the ever-born sibship size coefficients are qualitatively similar but somewhat smaller in magnitude.

The economic and demographic aggregates come from a variety of sources. I use cohort average outcomes from the DHS; GDP *per capita* and the sectoral composition of value added from the Penn World Table (Heston et al. 2012); average adult (ages 25+) educational attainment from Barro and Lee (2010) and Cohen and Soto (2007);[20] urbanization from UNPD (2011); and women's (ages 20-59) labor force participation from ILO (2012). For variables that are not available annually, I first linearly interpolate between observations within each country.

7.1 Cross-Sectional Patterns

Although the main analysis of economic and demographic aggregates takes advantage of the panel structure of the data by controlling for country and birth period fixed effects, cross-sectional analyses serve as a useful starting point. Figure 6 documents the evolution of cross-sectional relationships between several aggregate variables and γ_{ct}. Three of the four panels—for GDP *per capita*, average education, and urbanization—display a series of local linear regressions, one per period of birth. Data on women's labor force participation are too sparse to estimate cohort-level local linear regressions, so the fourth panel shows a scatter plot.

Throughout the sample period, more educated and more urban places have more negative sibsize-education associations. Although the intercepts shift downward over time, the slopes on these two curves are stable. These patterns suggest that structural transformation or mass education may be linked to the reversal of γ_{ct}. Meanwhile, γ_{ct} shows no consistent relationship with GDP *per capita* or women's labor force participation. The relationship between and log GDP *per capita* goes from flat to significantly negative, at least if one ignores the extreme outlier of Gabon.[21] No discernible pattern emerges in the scatter plot of γ_{ct} and women's labor force participation.

[20] I use the Barro-Lee estimates when available. For countries that only have Cohen-Soto estimates, I use the Cohen-Soto estimates to generate predicted Barro-Lee estimates, based on a regression of Barro-Lee on Cohen-Soto in the sample of countries with both measures.

[21] Gabon's oil production *per capita* is more than twice that of any other country in the sample, so its GDP *per capita* provides a poor measure of living standards.

Another noteworthy cross-sectional result, not reported in Figure 6, is that γ_{ct} in polygamous countries exceeds that in monogamous countries by 0.1 to 0.2, both within Africa and across the world. This finding supports Tertilt's (2005) claim that men in polygamous societies have an incentive to invest their wealth in a large number of children. In such societies, a groom typically 'buys' a bride from her father, so men benefit from having many daughters but do not lose from having many sons.[22]

7.2 Panel Analysis

The patterns in Figure 6 lead one to ask whether changes in socioeconomic and demographic aggregate can account for the reversal of the sibsize-education relationship. One can address this question by including cohort and country fixed effects:

$$\hat{\gamma}_{ct} = Z'_{ct}\lambda + \tau_t + \mu_c + \varepsilon_{ct} \tag{13}$$

where Z_{ct} is a vector of independent variables, and τ_t and μ_c are cohort and country fixed effects, respectively. This specification nets out global trends and time-invariant country characteristics. If one leaves Z_{ct} out of Equation (13), the resulting cohort effect estimates are flat through the early 1960s, at which point they begin a downward trend, becoming significantly negative in the 1970s. The estimates imply that net of country fixed effects, the sibsize-education association is 0.28 lower in 1985-9 than in 1945-9.

7.2.1 Using Cohort Average Outcomes as Covariates

Table 3 presents estimations of Equation (13) in which the covariates Z_{ct} are cohort average outcomes from the DHS: average completed education, average surviving sibship size, and the average fraction of siblings dying before they reach age 5. Because these average outcomes are co-determined with the sibsize-education relationship, one should think of the estimates equilibrium associations rather than causal effects. For this reason, I include only one covariate in each regression (in addition to the cohort and country fixed effects). Also, because the estimates of γ_{ct} and the

[22] Note that the patterns here must be driven by the number of children per wife, not the number of wives per husband. The DHS sibling history asks for siblings with the same biological mother.

cohort average outcomes are based on the same data, the table supplements the ordinary least squares results with estimations that correct for correlated measurement errors using Fuller's (1987) method-of-moments technique.

The results in Table 3 give three conclusions: (1) as the sibsize-education association declines, average educational investment increases, (2) as the sibsize-education association declines, average family size declines, and (3) the sibsize-education association has no relation to child mortality rates. These findings are consistent with explanations based on rising incomes, rising skill returns, and declining child labor, but not with those based on rising human capital endowments, declining child mortality, or an unmet need for contraception.

7.2.2 Using Socioeconomic Aggregates in Early Life as Covariates

Table 4 estimates regressions of $\hat{\gamma}_{ct}$ on three socioeconomic aggregates in the period of birth: log GDP *per capita*, average adult educational attainment, and urbanization. The education measure comes from two datasets that do not completely overlap, so the table presents one regression for the combined sample and one regression for each of the source samples. All three regressions lead to the same conclusion: while aggregate income growth and urbanization do not play a role, the rising educational attainment of the parent generation is intimately connected with the reversal of the sibsize-education relationship among offspring.[23] In fact, the coefficient of -0.1 on average education implies that rising education can account for roughly 60% of the of 1985-9 cohort effect effect for γ_{ct}, as reported at the start of this section. These results best match explanations based on rising skill returns or changing preferences, but if average education is better than GDP as a proxy for long-term parental income, then the results also support the income-gains hypothesis.

Several alternative theories deal with the position of women; these theories are the focus of Table 5.[24] One prominent theory involves the expansion of women's labor market opportunities outside the home. Recall that this explanation predicts a role for

[23]One could argue that average adult education violates the strict exogeneity assumption of fixed effects estimation because it is a function of the lagged sibsize-education relationship. First difference estimation, which does not assume strict exogeneity, yields a similar coefficient on adult education.

[24]To maximize sample size, each regression in Table 5 uses a different sample. In unreported results, average adult education has at least a marginally significant effect on γ_{ct} in each of these samples.

both rising women's labor force participation and the emergence of the service sector (which relocates women's work from near the home to far away). Columns (1) and (2) show that neither trend plays a role in the reversal of the sibsize-education association. Another gender-specific theory emphasizes female education over male. Column (3) thus uses gender-disaggregated data from the Barro-Lee education dataset to ask whether the role of average education is due to women or men.[25] While the coefficients on average female education and average male education are jointly significantly different from zero, they are not significantly different from each other; in fact, the coefficient on average male education is larger and individually more significant. Table 5 suggests that the causes of the reversal are not specific to the empowerment of women.

8 Conclusion

Prior to the results of this paper, limited evidence existed on positive associations between income and fertility or between sibship size and education in the 20^{th} century. The lack of extensive evidence led many researchers to focus instead on the negative associations widely observed today. A wide range of data from 48 developing countries reveals that both associations were indeed positive well into the 20^{th} century. They became negative only recently: first in Latin America, then in Asia, and finally in Africa. Increases in the aggregate education levels of the parents' generation were by far the most important predictor of the reversal; the data show little role for child mortality rates, GDP *per capita*, sectoral composition, urbanization, and women's labor force participation. Given the unique role of rising aggregate education and the leftward shift of the peak of the fertility-durable goods relationship, the data are most consistent with a theory in which a rising return to schooling leads families further and further down the income distribution to invest in education. As poorer families begin to invest in education, the relationship between income and fertility (and between durable goods ownership and fertility) turns negative for them.

Because the reversal has gone largely unrecognized in the literature on the aggregate effects of differential fertility, that literature has missed an important aspect of

[25] The Cohen-Soto education dataset does not provide gender-specific averages.

the interaction between demography and economic growth. In the mid-20th century, fertility differences by parental income increased average education in most of the countries under study. These fertility differences eventually flipped in many countries, so the effects of differential fertility on the *per capita* stock of human capital also reversed later in the century. A fruitful direction for future research would investigate the general equilibrium implications of these changes for the evolution of income inequality.

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Figure 1: Optimal Fertility and Education Spending

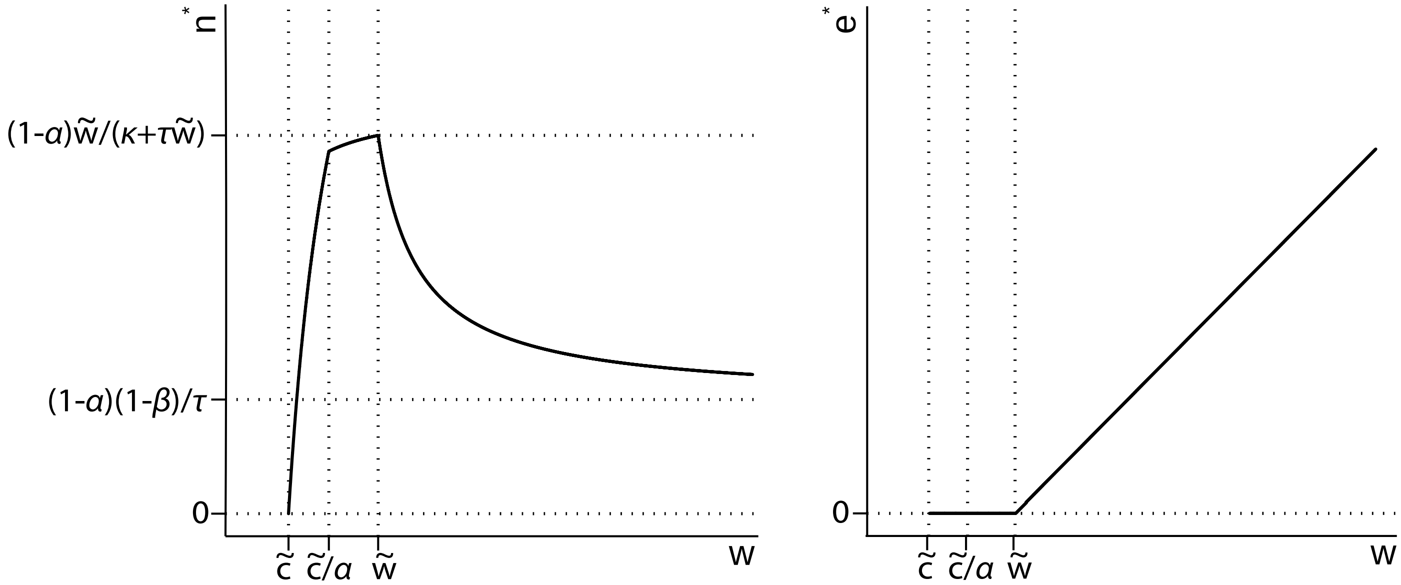
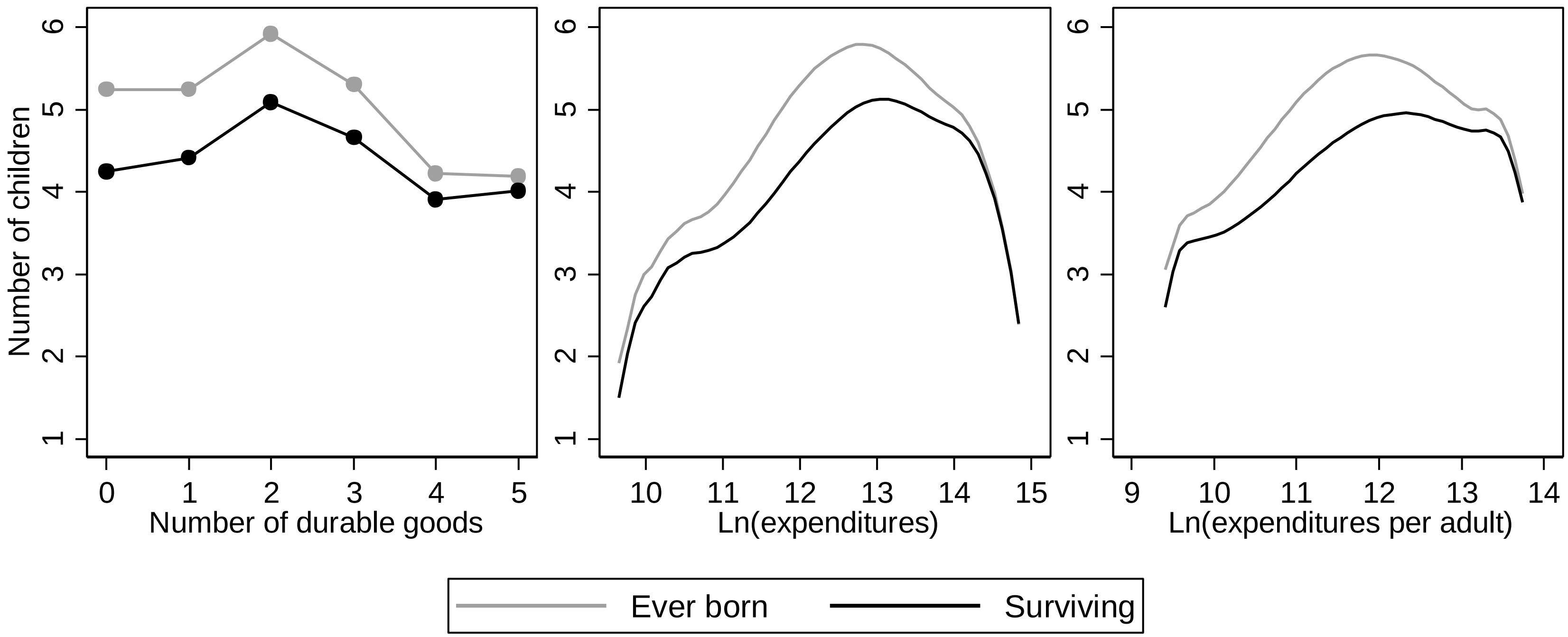
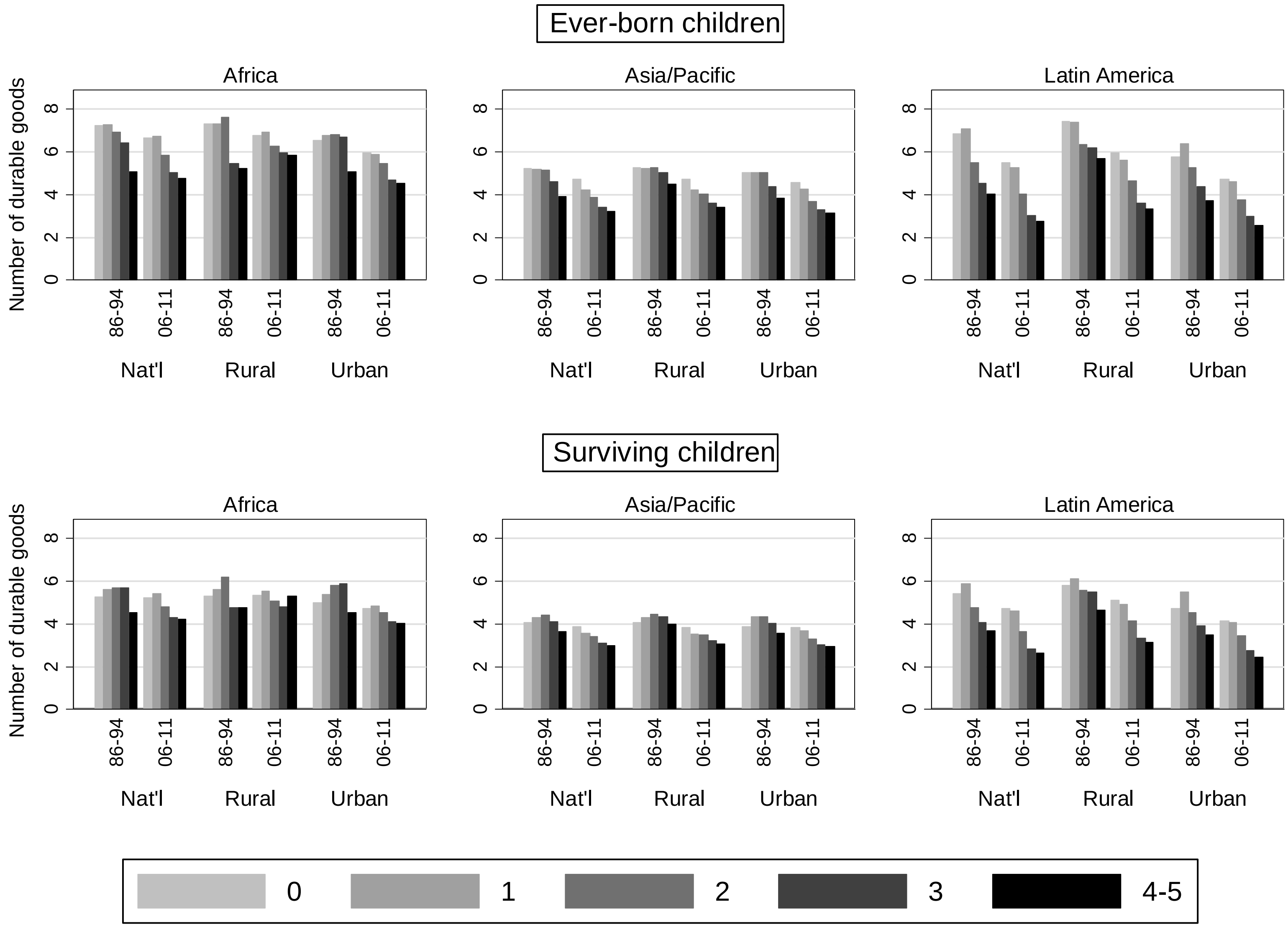


Figure 2: Completed Fertility and Three Measures of Household Consumption, Indonesia



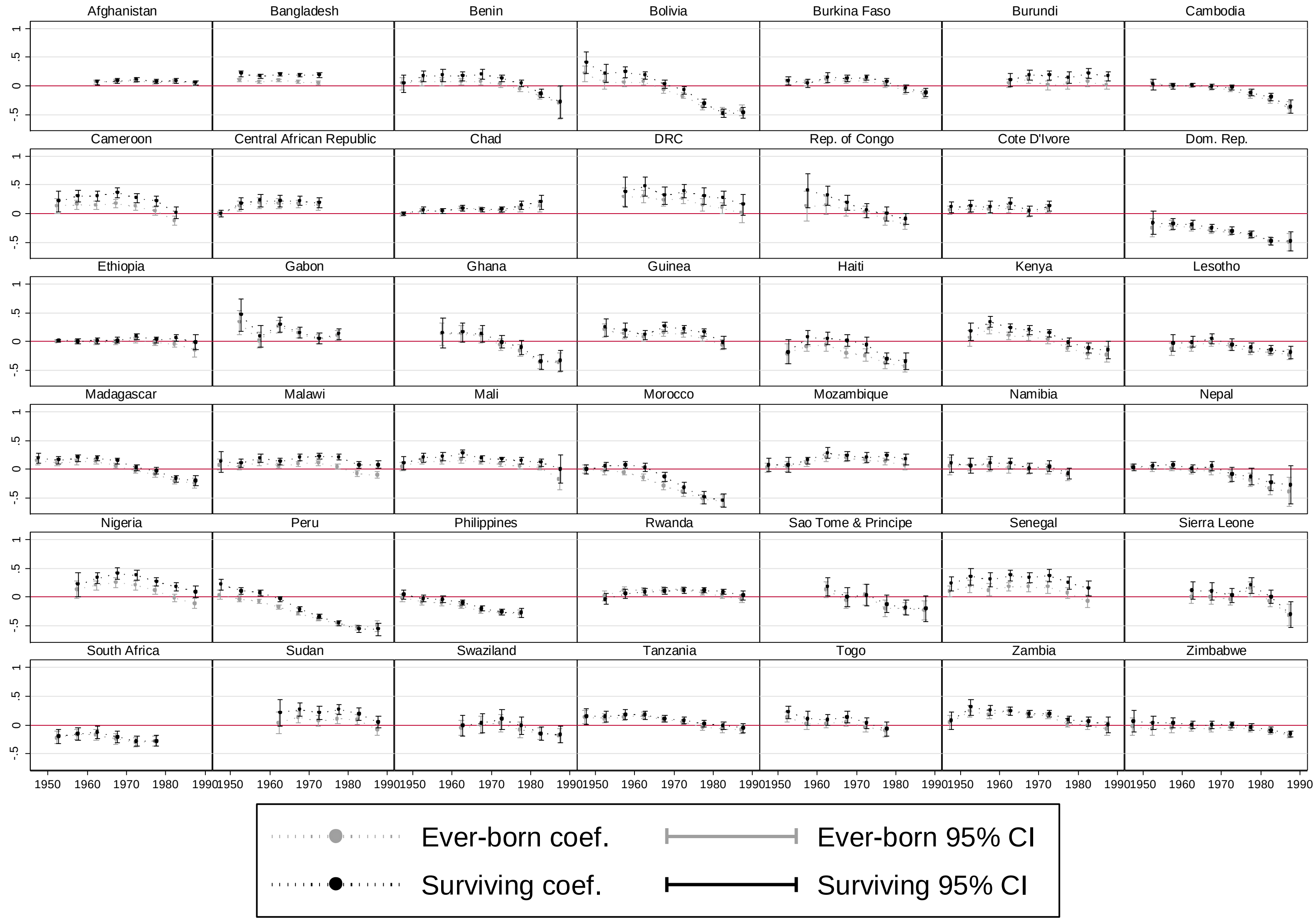
Note: The durables index is the sum of ownership dummies for radio, television, refrigerator, motorcycle, and car. Expenditures are measured in 1994 Rupiahs per month; adults are defined as household members over age 25. Data source: women age 45-49 in the Indonesia 1994 DHS Fertility History.

Figure 3: Completed Fertility by Number of Durable Goods Owned



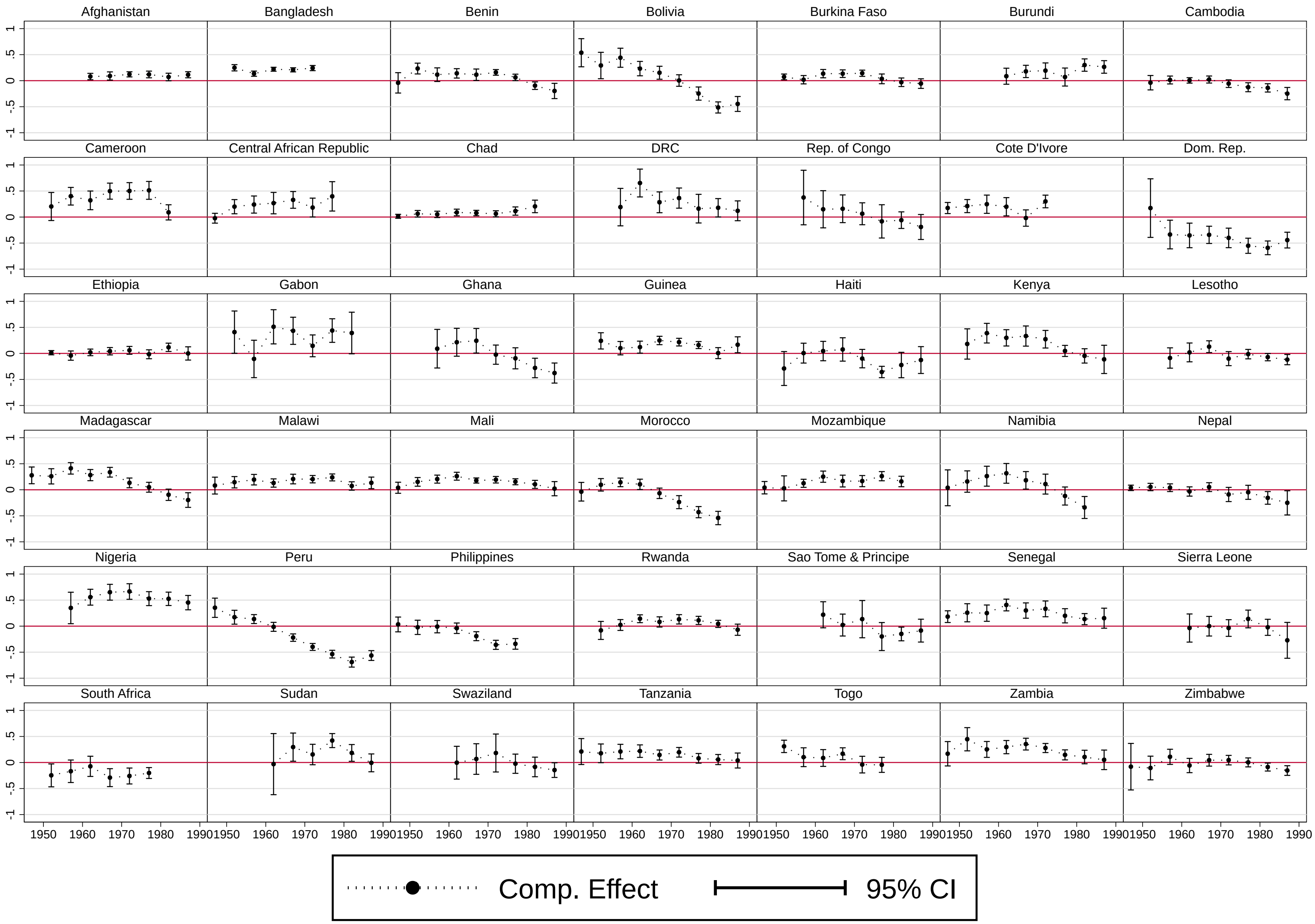
Note: Continental averages of country-specific averages. The durables index is the sum of ownership dummies for radio, television, refrigerator, motorcycle, and car. Data source: women age 45-49 in the DHS Fertility Histories.

Figure 4: Education-Sibship Size Coefficients by Period of Birth



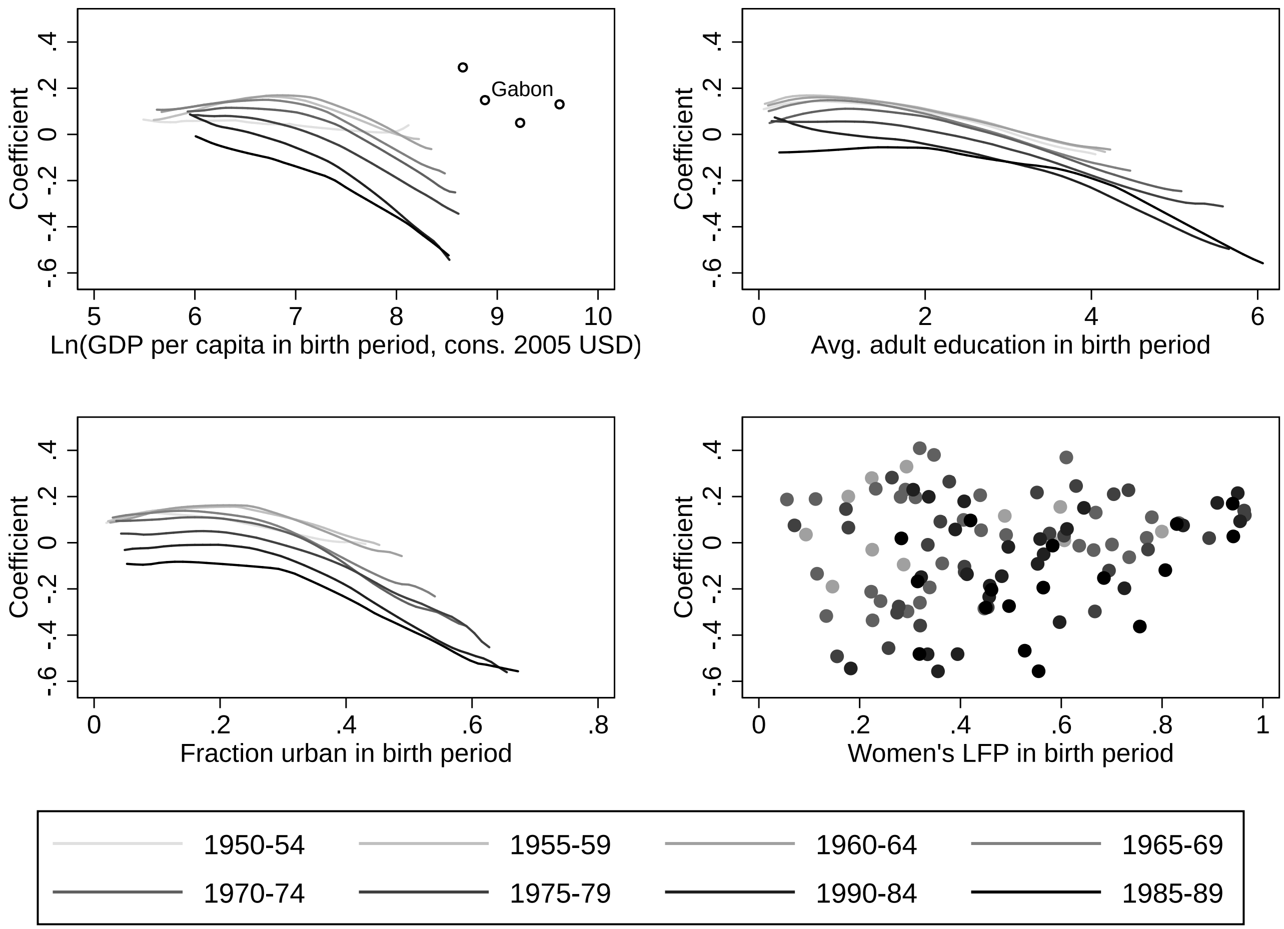
Note: From regressions of years of education on sibship size. Data source: DHS Sibling Histories.

Figure 5: Composition Effects of Differential Fertility by Period of Birth



Note: The composition effect is the difference between actual average education and the counterfactual that would arise if all families had the same number of siblings, with no change to their education. CIs are calculated with the delta method. Data source: DHS Sibling Histories.

Figure 6: Cross-Sectional Determinants of the Education-Sibship Size Relationship



Note: 307 observations from 42 countries. The dependent variable is the coefficient from a regression of education on surviving sibship size. Data source: DHS Sibling Histories.

Table 1: Household Durable Goods Ownership and Completed Fertility

	Africa (Burkina Faso, Burundi, Cameroon, Ghana, Kenya, Madagascar, Malawi, Namibia, Niger, Nigeria, Senegal, Tanzania, Zambia, Zimbabwe)				Asia/Pacific (India, Indonesia)				Latin America/Caribbean (Colombia, Dominican Republic, Haiti, Peru)			
	Ever-born		Surviving		Ever-born		Surviving		Ever-born		Surviving	
	'86-'94	'06-'11	'86-'94	'06-'11	'86-'94	'06-'11	'86-'94	'06-'11	'86-'94	'06-'11	'86-'94	'06-'11
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
A. Urban and rural areas												
A1. Without urban residence indicator												
Durables index	-0.316*	-0.565*†	0.095	-0.314*†	-0.236*	-0.368*†	-0.023	-0.212*†	-0.883*	-0.780*	-0.582*	-0.599*
	[0.066]	[0.037]	[0.056]	[0.031]	[0.033]	[0.022]	[0.028]	[0.019]	[0.058]	[0.030]	[0.048]	[0.028]
N	6,269	13,860	6,269	13,860	11,721	14,007	11,721	14,007	3,929	12,313	3,929	12,313
A2. With urban residence indicator												
Durables index	-0.176*	-0.353*†	0.158*	-0.165*†	-0.188*	-0.332*†	-0.002	-0.196*†	-0.637*	-0.611*	-0.403*	-0.466*
	[0.071]	[0.040]	[0.059]	[0.034]	[0.037]	[0.024]	[0.031]	[0.020]	[0.061]	[0.035]	[0.052]	[0.031]
Urban	-0.852*	-1.253*†	-0.433*	-0.878*†	-0.270*	-0.238*	-0.115	-0.092	-1.356*	-1.037*	-1.004*	-0.843*
	[0.140]	[0.093]	[0.114]	[0.080]	[0.109]	[0.084]	[0.092]	[0.071]	[0.181]	[0.087]	[0.152]	[0.079]
N	6,269	13,860	6,269	13,860	11,721	14,007	11,721	14,007	3,929	12,313	3,929	12,313
B. Urban Areas												
Durables index	-0.218*	-0.507*†	0.081	-0.278*†	-0.328*	-0.342*	-0.124*	-0.223*	-0.703*	-0.609*	-0.477*	-0.480*
	[0.084]	[0.047]	[0.073]	[0.040]	[0.045]	[0.035]	[0.042]	[0.029]	[0.068]	[0.042]	[0.060]	[0.037]
N	1,704	4,212	1,704	4,212	3,654	6,242	3,654	6,242	2,513	8,123	2,513	8,123
C. Rural Areas												
Durables index	-0.069	-0.229*	0.280*	-0.078†	-0.057	-0.324*†	0.112*	-0.175*†	-0.454*	-0.620*	-0.212*	-0.452*†
	[0.108]	[0.058]	[0.089]	[0.051]	[0.056]	[0.033]	[0.044]	[0.028]	[0.124]	[0.058]	[0.105]	[0.053]
N	4,565	9,648	4,565	9,648	8,067	7,765	8,067	7,765	1,416	4,190	1,416	4,190

Note: Each entry is an average of country-specific coefficients; standard errors are in brackets. The durables index is the sum of ownership dummies for radio, television, refrigerator, motorcycle, and car. Each regression controls for age and survey year indicators, and clusters standard errors at the PSU level. Sample sizes are the sum of the country-specific sample sizes. The sample includes a country if and only if it had at least one standard DHS survey with a full durable goods module in both the early and late periods. * sig. diff. from zero at 5% level; † sig. diff. from the early-period coefficient at 5% level. Data source: women age 45-49 in the DHS Fertility Histories.

Table 2: Demographic Correlates of the Education-Sibship Size Relationship

	Mean (SD)	OLS	Fuller	OLS	Fuller	OLS	Fuller
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Cohort average education	4.2	-0.045	-0.051				
	(2.8)	[0.021]**	[0.024]**				
Cohort average surviving	4.4			0.090	0.096		
sibship size	(0.7)			[0.036]**	[0.047]**		
Cohort average fraction of	0.10					0.42	0.63
siblings dying under 5	(0.04)					[0.82]	[1.52]
Number of observations	307	307	307	307	307	307	307
Number of countries	42	42	42	42	42	42	42
Birth Cohort and Country FE		X	X	X	X	X	X

Note: The dependent variable is the coefficient from a regression of education on surviving sibship size. Brackets contain standard errors clustered at the country level. The Fuller GMM estimates are block-bootstrapped. * sig. at the 10% level; ** sig. at the 5% level. Data source: DHS Sibling Histories.

Table 3: Development and the Education-Sibship Size Relationship

	OLS	OLS	OLS
	(1)	(2)	(3)
Ln(GDP per capita in birth period)	0.038	0.022	0.012
	[0.091]	[0.094]	[0.094]
Avg. adult yrs. ed. in birth period	-0.093	-0.106	-0.110
	[0.025]**	[0.029]**	[0.032]**
Fraction urban in birth period	-0.633	-0.556	-0.284
	[0.441]	[0.468]	[0.393[
Number of observations	217	193	142
Number of countries	38	34	27
Education dataset	Combined	Barro-Lee	Cohen-Soto
Birth Cohort and Country FE	X	X	X

Note: The dependent variable is the coefficient from a regression of education on surviving sibship size. Brackets contain standard errors clustered at the country level. * sig. at the 10% level; * sig. at the 5% level.

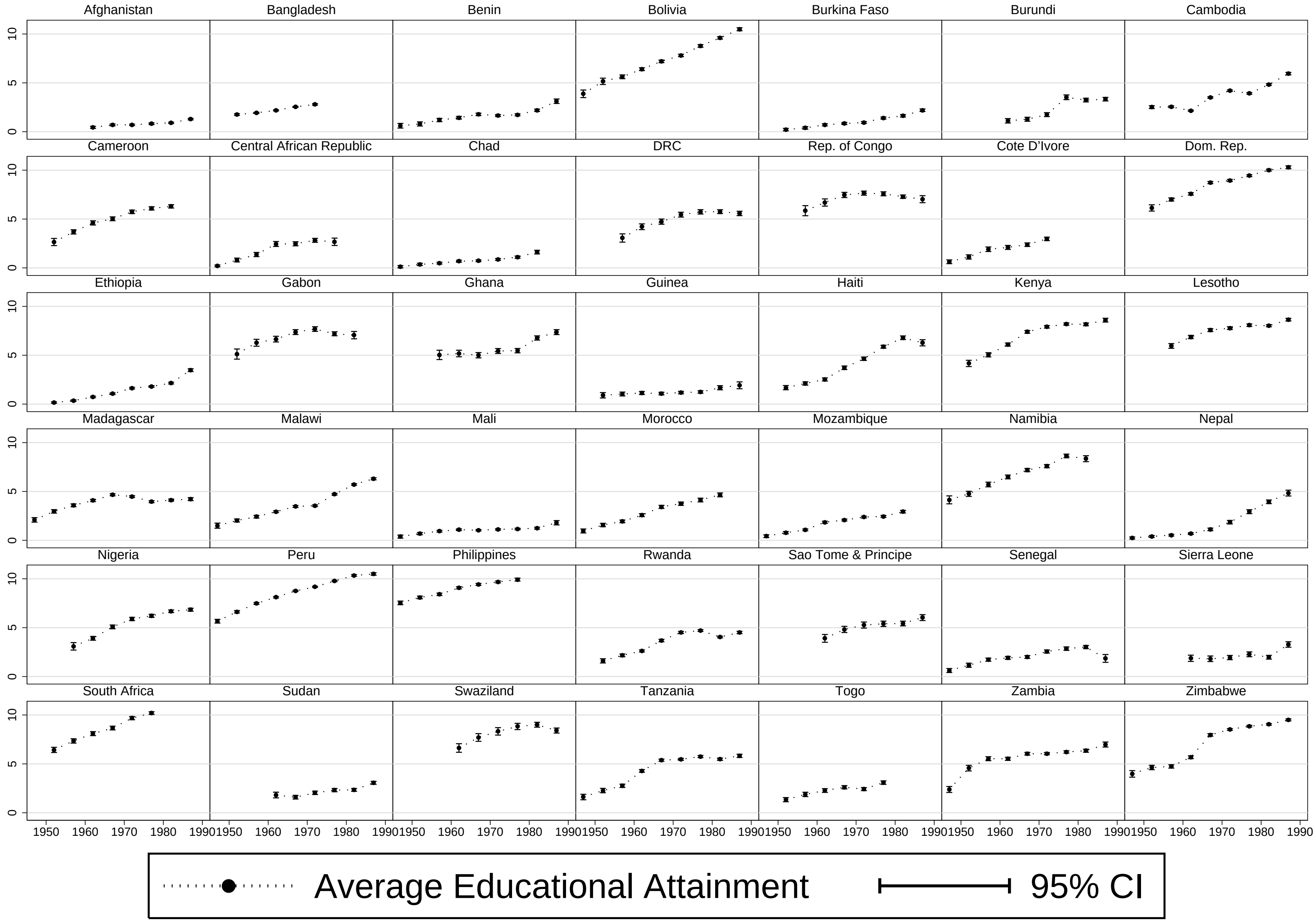
Table 4: Female Empowerment and the Education-Sibship Size Relationship

	OLS	OLS	OLS
	(1)	(2)	(3)
Women's labor force participation rate in birth period	0.113		
	[0.111]		
Manufacturing fraction of value added in birth period		-0.001	
		[0.002]	
Services fraction of value added in birth period		0.001	
		[0.002]	
Avg. adult male yrs. ed. in birth period			-0.065
			[0.022]**
Avg. adult female yrs. ed. in birth period			-0.056
			[0.037]
p-value: joint test of education coefficients			0.002
p-value: difference of education coefficients			0.851
Number of observations	112	137	234
Number of countries	34	41	34
Birth Cohort FE	X	X	X
Country FE	X	X	X

Note: The dependent variable is the coefficient from a regression of education on surviving sibship size. Brackets contain standard errors clustered at the country level. The education averages in column (3) are from the Barro-Lee dataset. * sig. at the 10% level; * sig. at the 1% level.

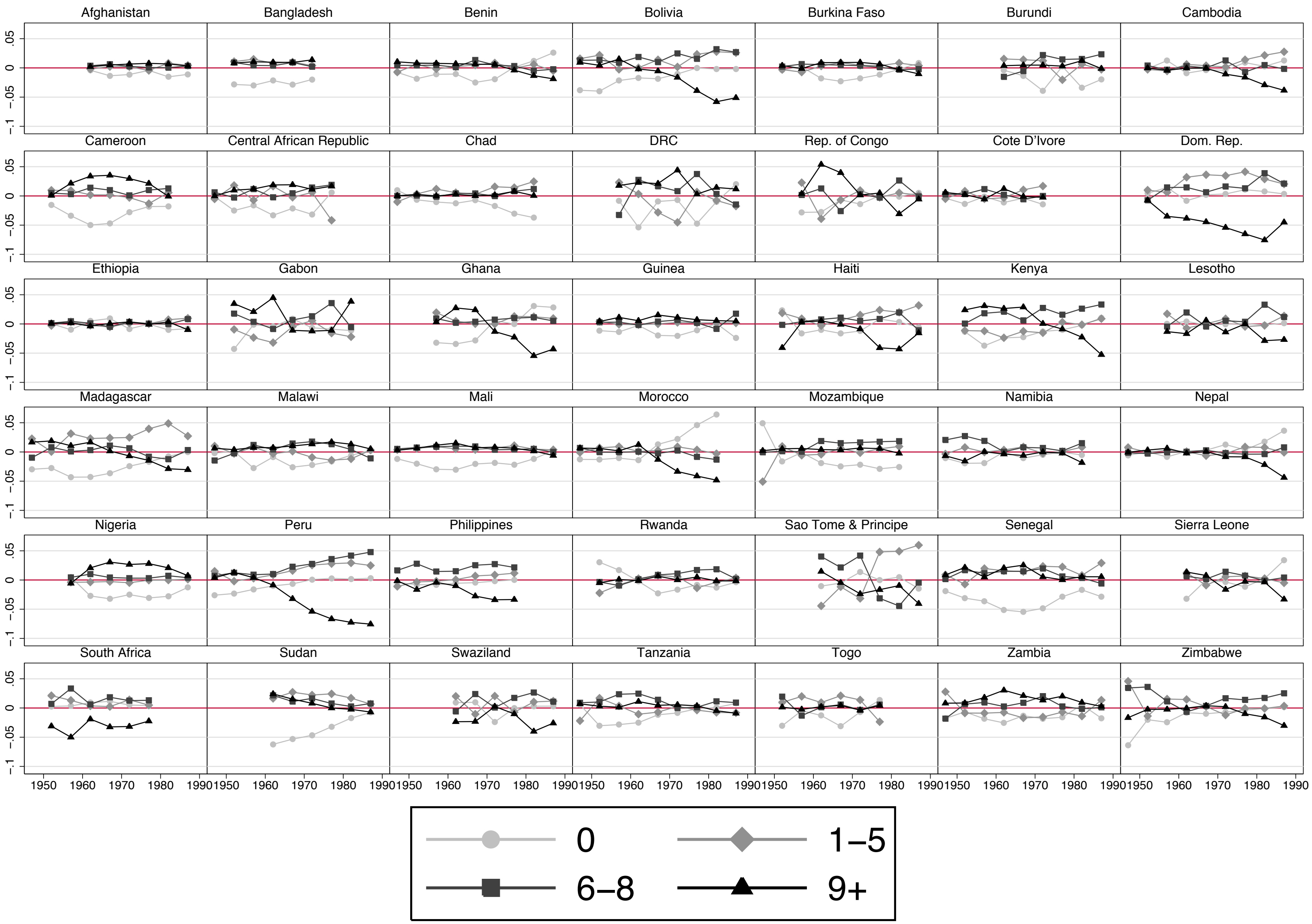
Appendix Tables and Figures — Not for Publication

Appendix Figure 1: Mean Education by Period of Birth



Data source: DHS Sibling Histories.

Appendix Figure 2: Composition Effects on Education Shares by Period of Birth



Note: The composition effect is the difference between actual average education and the counterfactual that would arise if all families had the same number of siblings, with no change to their education. Data source: DHS Sibling Histories.

Appendix Table 1: Avgs. of Country-Specific Means and Standard Deviations in the Fertility Histories

	Africa (Burkina Faso, Burundi, Cameroon, Ghana, Kenya, Madagascar, Malawi, Namibia, Niger, Nigeria, Senegal, Tanzania, Zambia, Zimbabwe)		Asia/Pacific (India, Indonesia)		Latin America/Caribbean (Colombia, Dominican Republic, Haiti, Peru)	
	‘86-'94	‘06-'11	‘86-'94	‘06-'11	‘86-'94	‘06-'11
	(1)	(2)	(3)	(4)	(5)	(6)
Ever-born fertility	7.14	6.32	5.12	4.02	5.79	4.04
	[3.22]	[2.88]	[2.67]	[2.29]	[3.43]	[2.44]
Surviving fertility	5.39	5.12	4.19	3.45	4.89	3.59
	[2.69]	[2.49]	[2.22]	[1.90]	[2.89]	[2.16]
Durables index	0.68	1.26	1.12	1.77	1.79	2.24
	[0.86]	[1.10]	[1.22]	[1.38]	[1.22]	[1.02]
Woman’s years of education	1.39	3.60	2.88	4.21	3.63	6.31
	[2.27]	[3.76]	[3.75]	[4.41]	[3.64]	[4.59]
Husband’s years of education	2.52	4.57	4.88	6.26	4.65	7.31
	[3.03]	[4.21]	[4.68]	[4.78]	[4.25]	[4.88]
Urban	0.21	0.30	0.28	0.38	0.57	0.65
	[0.39]	[0.43]	[0.45]	[0.48]	[0.47]	[0.44]
N	6,269	13,860	11,721	14,007	3,929	12,313

Note: Average means, with average standard deviations in brackets. Each entry represents a simple average of country-specific statistics. The mean for husband’s years of education is for the subsample with non-missing values on that variable (roughly 94% of the overall sample). Sample sizes refer to the sum of the country-specific sample sizes. The sample includes a country if and only if it was the site of at least one standard DHS survey with a full durable goods module in both the early and late periods. The durables index is the sum of ownership dummies for the following durable goods: radio, television, refrigerator, motorcycle, and car. Data source: women age 45-49 in the DHS Fertility Histories.

Appendix Table 2: Socioeconomic Characteristics and Completed Fertility

	Africa (Burkina Faso, Burundi, Cameroon, Ghana, Kenya, Madagascar, Malawi, Namibia, Niger, Nigeria, Senegal, Tanzania, Zambia, Zimbabwe)				**Asia/Pacific** (India, Indonesia)				**Latin America/Caribbean** (Colombia, Dominican Republic, Haiti, Peru)			
	Ever-born		Surviving		Ever-born		Surviving		Ever-born		Surviving	
	‘86-'94	‘06-'11	‘86-'94	‘06-'11	‘86-'94	‘06-'11	‘86-'94	‘06-'11	‘86-'94	‘06-'11	‘86-'94	‘06-'11
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Durables index	-0.120*	-0.070	0.152*	-.0002†	0.007	-0.145*†	0.079*	-0.078*†	-0.391*	-0.293*	-0.234*	-0.212*
	[0.075]	[0.041]	[0.064]	[0.036]	[0.044]	[0.025]	[0.037]	[0.022]	[0.064]	[0.038]	[0.057]	[0.035]
Woman’s years of education	-0.067	-0.156*†	-0.044	-0.102*	-0.098*	-0.109*	-0.063*	-0.079*	-0.140*	-0.147*	-0.100*	-0.118*
	[0.038]	[0.013]	[0.036]	[0.012]	[0.015]	[0.010]	[0.013]	[0.008]	[0.021]	[0.010]	[0.018]	[0.009]
Husband’s years of education	0.026	-0.061*†	0.059*	-0.025*†	-0.025	-0.017	0.008	-0.002	-0.102*	-0.059*	-0.075*	-0.047*
	[0.029]	[0.011]	[0.025]	[0.010]	[0.014]	[0.010]	[0.012]	[0.008]	[0.020]	[0.011]	[0.018]	[0.010]
Urban	-0.795*	-0.815*	-0.431*	-0.603*	-0.064	-0.012	-0.029	0.063	-0.852*	-0.489*	-0.621*	-0.392
	[0.143]	[0.090]	[0.118]	[0.079]	[0.111]	[0.08]	[0.096]	[0.071]	[0.178]	[0.079]	[0.147]	[0.072]
N	6,269	13,860	6,269	13,860	11,721	14,007	11,721	14,007	3,929	12,313	3,929	12,313

Note: Each entry represents a simple average of country-specific coefficients, with the associated standard error in brackets. The durables index is the sum of ownership dummies for the following durable goods: radio, television, refrigerator, motorcycle, and car. Each country-specific regression controls for single-year age indicators and survey year indicators, and clusters standard errors at the PSU level. Sample sizes refer to the sum of the country-specific sample sizes. The sample includes a country if and only if it was the site of at least one standard DHS survey with a full durable goods module in both the early and late periods. * sig. diff. from zero at the 5% level; † sig. diff. from the early-period coefficient at the 5% level. Data source: women age 45-49 in the DHS Fertility Histories.

Appendix Table 3: Education-Sibship Size Coefficients by Gender and Period of Birth

	1940-1949	1950-1959	1960-1969	1970-1982
	(1)	(2)	(3)	(4)
Indonesia				
Men	0.399	0.427	0.303	0.179
	[0.074]**	[0.063]**	[0.070]**	[0.156]
N	949	1,450	1,133	132
Women	0.418	0.383	0.295	0.085
	[0.065]**	[0.046]**	[0.056]**	[0.098]
N	1,076	1,614	1,762	479
Matlab, Bangladesh				
Men	0.309	0.274	0.172	0.143
	[0.086]**	[0.070]**	[0.076]*	[0.077]
N	751	920	894	780
Women	0.123	0.249	0.141	0.062
	[0.028]**	[0.039]**	[0.039]**	[0.067]
N	968	1,130	1,481	967
Mexico				
Men	0.05	-0.023	-0.186	-0.29
	[0.088]	[0.086]	[0.067]**	[0.045]**
N	845	1,256	1,644	2,154
Women	0.017	-0.038	-0.127	-0.29
	[0.066]	[0.068]	[0.052]*	[0.044]**
N	966	1,574	2,222	3,053

Note: OLS coefficients. Brackets contain standard errors clustered at the PSU level. Each coefficient is from a separate regression. * different from zero at 5% level; ** different from zero at 1% level. Data source: adults born between 1940 and 1982 in the Indonesia Family Life Survey (1993, 1997 waves), Matlab Health and Socioeconomic Survey (1996), and Mexico Family Life Survey (2002 wave).

Appendix Table 4: Education-Sibship Size Coefficients with and without Controlling for Father's Education

	1940-1949	1950-1959	1960-1969	1970-1982
	(1)	(2)	(3)	(4)
Indonesia				
Unadjusted for dad's ed.	0.344	0.415	0.328	0.074
	[0.061]**	[0.051]**	[0.058]**	[0.103]
Adjusted for dad's ed.	0.189	0.225	0.119	-0.012
	[0.056]**	[0.040]**	[0.043]**	[0.088]
N	1,430	2,049	2,009	460
Matlab, Bangladesh				
Unadjusted for dad's ed.	0.191	0.264	0.160	0.093
	[0.040]**	[0.038]**	[0.037]**	[0.052]
Adjusted for dad's ed.	0.102	0.138	0.071	0.119
	[0.037]**	[0.036]**	[0.034]*	[0.046]**
N	1,678	2,007	2,317	1,705
Mexico				
Unadjusted for dad's ed.	0.032	-0.037	-0.162	-0.301
	[0.080]	[0.067]	[0.057]**	[0.037]**
Adjusted for dad's ed.	0.071	0.0002	-0.045	-0.154
	[0.070]	[0.057]	[0.048]	[0.034]**
N	1,376	2,261	3,166	4,393

Note: OLS coefficients. Brackets contain standard errors clustered at the PSU level. Each coefficient is from a separate regression. The samples include both men and women, and all regressions control for a gender indicator. * different from zero at 5% level; ** different from zero at 1% level. Data source: adults born between 1940 and 1982 in the Indonesia Family Life Survey (1993, 1997 waves), Matlab Health and Socioeconomic Survey (1996), and Mexico Family Life Survey (2002 wave). The Mexico Family Life Survey only contains data on broad education categories, but for ease of comparison across settings, I convert them to a measure of years of education. Using data from the 2000 Mexico census, I determine the mean years of education among adults in each education level, and I then assign that mean to the corresponding parents in the Mexico Family Life Survey.