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Fertility Decline, Women's Well-being, and Gender Gaps in Well-being in Poor Countries

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Fertility Decline, Women's Well-being, and Gender Gaps in Well-being in Poor Countries

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ABSTRACT

We examined how declining total fertility and women's increasing median age at first birth have been associated with changes in women's and girls' well-being and gender gaps in children's well-being in panels of 60–75 poorer countries using 124–187 Demographic and Health Surveys spanning 1985–2008. In adjusted random-effects models, these changes in fertility were associated with gains in women's and girls' well-being, particularly access to prenatal care and trained attendance at delivery, survival at 1–4 years, vaccination coverage at 12–23 months, school attendance at 11–15 years, and nutrition at 0–36 months (for later childbearing). Benefits were equal for boys and girls with respect to vaccination coverage and school attendance. Declining total fertility was associated with greater gains for boys relative to girls with regards to child mortality and malnutrition; however, increases in women's age at first birth were associated with greater advantages for girls relative to boys on these same measures. Family planning programs in higher-fertility societies may wish to encourage equitable investments in children.

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INTRODUCTION

The second half of the 20th century witnessed dramatic declines in fertility, especially in poorer countries (Bongaarts 2008), as well as increasing ages at first birth (Frejka, Jones and Sardon 2010; Frejka and Sardon 2006; Westoff 2003).[1] In the 1950s, for example, the median person in the world's population was living in a country in which total fertility was 5.4 births per woman of reproductive age; by 2000, the median person was living in a country in which total fertility was 2.3 (Wilson 2001). Also, in 27 countries in Europe and the U.S., the mean age at first birth rose from 22.0 – 24.0 for cohorts born in 1940 to 22.1 – 28.4 for cohorts born in 1965 (Frejka and Sardon 2006). Among 47 lower-income countries elsewhere, 25 had a majority of women in the 1950 – 4 birth cohort who had a first birth before age 20; however, only 16 of 47 countries had such a majority in the 1975 – 9 birth cohort (14 of these 16 countries were in Sub-Saharan Africa) (Westoff 2003).

Demographers, for decades, have studied the reasons for these shifts in fertility (e.g., Bryant 2007; Hill 1992; Mauldin 1978, 1982; Mauldin and Segal 1988; Tsui and Bogue 1978), including changes in women's opportunities and gender relations (e.g., Adamchack and Ntseane 1992; Folbre 1983; London 1992; Malhotra, Vanneman and Kishor 1995; Mason 1984, 1987; Sanderson and Dubrow 2000; Weinberger, Lloyd and Blanc 1989). Yet, recent dramatic shifts in fertility regimes may have had marked influences on many aspects of women's lives (Cain 1993; Das Gupta and Bhat 1997; Das Gupta and Shuzhuo 1999; Govindasamy and Malhotra 1996; Mason 1997; McDonald 2000), including their survival, autonomy and power within the family, non-familial economic opportunities, and other aspects of their well-being. Surprisingly, these reciprocal influences have been understudied, especially in poorer countries.[2]

Here, we examined for over a 23-year period how declining total fertility and women's increasing median age at first birth were associated with changing investments in women's and girls' well-being, as well as changing gender gaps in these investments in poorer countries. Underpinning the analyses were three sets of hypotheses, that *declining fertility and later first childbearing will be*

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[1] Dorius (2008) reveals, however persuasively, that trends toward convergence in total fertility between rich and poor countries began only around 1995.

[2] See Mason (1997), Huber (1991), McDonald (2000) and Brewster and Rindfuss (2000) for important theoretical discussions of these reciprocal influences as well as Engelhardt, Kögel, and Prskawetz (2004), Engelhardt and Prskawetz (2004), Kögel (2004), and Matysiak and Vignoli (2008) for empirical investigations in wealthier Western countries.

associated in the aggregate with (1a) *improvements in women's basic needs for well-being, as reflected by reductions in under-nutrition,* (1b) *greater investments in women's intermediate needs for well-being, as reflected in their access to prenatal and trained delivery care,* (2a) *improvements in girls' basic needs for well-being, as reflected in their survival and nutritional status,* (2b) *greater investments in girls' intermediate needs for well-being, as reflected in their vaccination coverage and schooling attendance,* and (3) *variable influences on gender gaps in investments in children because of regional variations in the gender normative environment* (Das Gupta and Bhat 1997; Das Gupta and Shuzhuo 1999).

We tested our hypotheses by leveraging data from as many as 187 national Demographic and Health Surveys (DHS) undertaken in as many as 75 poorer countries from 1985 to 2008 (Measure DHS 2011). This source provided comparable, national time-series measures of total fertility and women's median age at first birth, investments in women's and girls' needs for well-being, and gender gaps in these investments in children. We linked these data with other sources of information on changes in national socioeconomic conditions for the same period. We then estimated and compared country-time fixed-effects and random-effects models to examine (a) how changes in national fertility regimes were associated with changing investments in women's and girls' needs for well-being and gender gaps in these investments in children and (b) the extent to which concurrent changes in socioeconomic conditions accounted for and conditioned these associations (e.g., Brewster and Rindfuss 2000). We tested the consistency of our findings across regions and baseline national attributes, as well as the robustness of our findings with alternative samples that compensated for country attrition in the DHS.

In what follows, we first define the components of well-being that we analyzed and describe the theoretical bases for our hypotheses. We then describe the sample, data sources, variables, and methods. Third, we present the results, including (a) univariate statistics and trends in our main variables; and (b) estimates from country-time fixed-effects and random-effects panel regressions of our explanatory variables and outcomes, unadjusted and adjusted for changing national socioeconomic conditions. Lastly, we discuss the implications of our findings for theory, research, and policy.

HUMAN NEEDS, THEIR SATISFACTION, AND WOMEN'S WELL-BEING

Doyal and Gough's (1991) *Theory of Human Need* informs our conceptualization of women's and girls' *well-being* and the *basic* and *intermediate human needs* that are foundational to it. *Well-being* refers to "a state of being with others, where human needs are met, where one can act meaningfully to pursue one's goals, and where one enjoys a satisfactory quality of life" (Economic and Social Research Council, Research Group on Wellbeing in Developing Countries 2007:1). Well-being, thus, is enabled through investments in certain *universal human needs*, or common "preconditions" that, if met, avoid serious objective harm and enable successful social participation (Gough, McGregor and Camfield 2007). These preconditions include most essentially *survival*, and beyond that, some measure of *physical health*. They also include *autonomy of agency*, or the "capacity to make informed choices about what should be done and how to go about doing it" (Gough et al. 2007:14). Autonomy of agency can be constrained by mental or cognitive impairments and by blocked opportunities to engage in social participation. The needs for survival, health, and autonomy of agency arguably can be satisfied by investing in and meeting a range of *intermediate needs*, which have material foundations or are pursued through social relationships. These intermediate needs include adequate nutrition, water, and housing; non-hazardous work and physical environments; appropriate health care; security in childhood; significant primary relationships; physical and economic security; and appropriate basic and cross-cultural education. The outcomes in this paper capture (a) changing investments in selected needs for women's and girls' well-being, (b) the changing burden of harms to women and girls from unmet needs, and (c) changing gender gaps in these investments and harms among children. Hereafter, we refer to these outcomes as measures of women's and girls' well-being and gender gaps in children's well-being.

TOTAL FERTILITY, AGE AT FIRST BIRTH, AND WOMEN'S AND GIRLS' WELL-BEING

Our main explanatory variables focus on declines in total fertility and rising median ages at first birth for women because these variables capture different aspects (and often stages) of fertility transition. Namely, the total fertility rate (TFR) may decline with an increase in women's median age at first birth[3] and with restrictions in fertility at older ages, both of which shorten women's average reproductive life span. In Asia, Europe, and Latin America (Andersonn et al. 2009; Knodel 1977; Rosero-Bixby, Castro-Martín and Martín-García 2009), restriction in fertility at older ages (e.g., 35 years and older) characterized initial declines in fertility, and later increases in the median age at first birth arguably have reflected changing gender norms toward greater equity in roles and

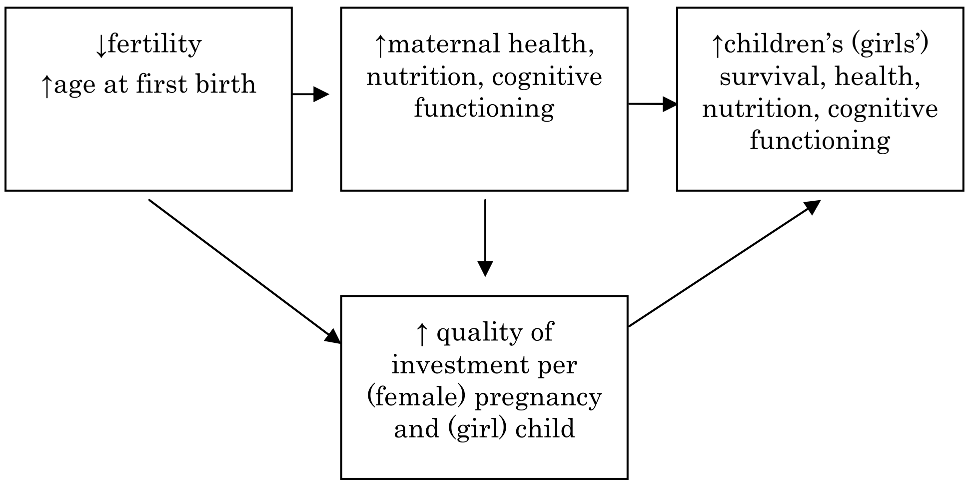
[3] Unless there are changes in the tempo of childbearing after first birth (Andersson et al. 2009).

opportunities (e.g., Andersonn et al. 2009; Presser 1971; Rosero-Bixby et al. 2009).[4] Although early fertility declines in Africa were first described as uniform across women's reproductive life span (Caldwell, Orubuloye and Caldwell 1992), analyses of 74 DHS in 30 sub-Saharan African countries between 1986 and 2006 have suggested that "fertility decline [also] began with women later in their reproductive years (25 years and older), followed by younger age groups" (Sneeringer 2009:21).

With these definitions in mind, *how* might declining total fertility and women's increasing median age at first birth enhance well-being among women and girls? To answer this question, our conceptual argument has two parts. First, such changes in fertility regimes should diminish the aggregate *reproductive stress* on women of frequent reproductive cycling, reducing the prevalence of maternal malnutrition (Merchant and Martorell 1988). Improved maternal health and nutrition should directly improve outcomes at birth and in early childhood for all children (King 2003). Second, the above changes in fertility regimes should create a normative environment that favors investments in pregnant mothers and children (Becker 1960; Becker and Lewis 1973; Swidler 1986). In particular, declining total fertility and later first childbearing should encourage the trade-off from a higher quantity to a higher quality of children (Becker 1960; Becker and Lewis 1973), and girls should benefit to some extent from this tradeoff (Figure 1). Below, we elaborate on these two arguments.

[4] McDonald (2000, p. 431) has argued that "implicit in the gender system of a high-fertility society is that women devote a great deal of their time and energy to childbearing and childrearing. If fertility falls to lower and lower levels, this in itself is an indication that society no longer places the same emphasis upon this division of labor." He argues further that the impact of fertility decline on women's lives may not occur in the early stages of fertility transition, but rather after a sustained period of decline, and correspondingly, a sustained expectation of improvement in women's lives.

Figure 1: Hypothesized Effects of Lower Fertility and Higher Ages at First Birth on Women's and Girls' Well-Being



NOTE. Adapted from Becker (1960), Becker & Lewis (1973)

Lower Fertility and Later First Childbearing Improve Women's Health and Nutrition

The first part of our argument focuses on the aggregate nutritional and physiological benefits to mothers of lower total fertility and later ages at first birth. Under such conditions, populations of women, on average, should experience less reproductive stress, both from child birth and child rearing, and should be better nourished and less often underweight (Koski-Rahikkala et al. 2006; Rah et al. 2008). Although some individual-level studies in energy-rich environments have shown that higher parity[5] is associated with higher crude and aged-adjusted body mass (e.g., Nenko and Jasienska 2009), other reviews and individual-level studies have shown that a later first birth is associated with lower risks of maternal mortality, morbidity, and nutritional depletion[6] (see King 2003; Rah et al. 2008). Pregnant adolescents in Bangladesh, for example, have had lower weight, BMI, mid-upper arm circumference, upper-arm muscle area, and percent body fat six months post-

[5] Or the number of children previously born alive to women.

[6] Maternal nutritional depletion is variously defined, but, in King (2003), the term refers to "a negative change in maternal nutritional status during a reproductive cycle going from nonpregnant, nonlactating to pregnancy; to lactation; to partial lactation; and back to nonpregnant, nonlactating" (p. 1734S, based on the definition in Winkvist, Rasmussen, and Habicht [1992]).

partum than their never-pregnant counterparts, suggesting depletion in energy stores resulting from pregnancy at a young age (Rah et al. 2008). In retrospective and prospective cohort studies, higher parity has been negatively associated with post-reproductive survival (Gagnon et al 2009; Smith, Mineau and Bean 2002) and positively associated with the adjusted risks of mortality from vascular complications (Koski-Rahikkala et al. 2006). In macro-level studies in South Asia, 35% of the 121,000 maternal lives saved between 1990 and 2008 have been attributed to declines in fertility in the region (Jain 2011). Thus, we argue that populations of women will experience, on average, less maternal depletion and will less often be underweight as the total fertility rate declines and the median age at first birth increases. Healthier and better nourished mothers should have, on average, better nourished newborn boys *and* girls who also are more likely to survive (e.g., Imdad, Yakoob and Bhutta 2011; Haider, Yakoob and Bhutta 2011).

Lower Fertility and Later First Childbearing Encourage Investments in Women and Girls

The second part of our argument adapts theories originated by Becker and colleagues of the quantity–quality tradeoff that occurs with fertility decline (Becker 1960; Becker and Lewis 1973). Specifically, *child quality* can be defined as a function of the time and resources that parents devote to each child (Willis 1973). With fewer children, families can invest more time and resources into each pregnancy and birth (Becker 1960; Becker and Lewis 1973; Hanushek 1992), including female fetuses and girls. In the aggregate, an environment characterized, on average, by smaller completed family sizes[7] and later ages at first birth should encourage families to invest more in each pregnancy and child. These new investments should meet a higher share of mothers' and children's needs for well-being. Evidence of the influences of smaller completed family size and later first childbearing are available at different levels of analysis for various settings and aspects of well-being.

Studies of parity and women's use of prenatal and delivery care suggest that lower fertility leads families to invest more in each pregnancy. Specifically, *lower* parity has been associated with more frequent prenatal care in parts of Kenya (Delva et al. 2010), along with higher odds of more than four prenatal visits, of delivering at a hospital, and of receiving postnatal care in rural western China (Liu et al. 2011). Conversely, in parts of Africa and South Asia, *higher* parity (variously

[7] By completed family size, we mean the total number of children that a woman has in her reproductive lifetime. We use "family size" and "completed family size" interchangeably.

defined) has been associated with lower uptake of: antenatal care (Doctor 2011; Overbosch et al. 2004), supervised deliveries (Doctor 2011), use of health professionals (Mills et al. 2008), and delivery at a health facility (Agha and Carton 2011; Ekele and Tunau 2007), as well as greater use of lower-cost prenatal care providers (Overbosch et al. 2004). In recent systematic and other reviews for poorer countries, higher parity has been significantly negatively associated with adequate prenatal care (Simkhada et al. 2008), and coverage of skilled care at birth decreases with increasing parity (Stanton et al. 2006). Finally, mothers below age 18 years have had lower odds than their counterparts of receiving prenatal and delivery care in Asia (Reynolds, Wong and Tucker 2006).

Other studies in wealthier and poorer countries suggest that lower fertility is associated with better health outcomes in children. Among Columbia-Medical-Plan enrollees, three-child families have had lower mean health-care use per family member than have one- or two-child families (Schor et al. 1987). Also in the U.S., higher birth-order children have been less likely than first births to have ever been breastfed (Hirschman and Butler 1981), and large family size has been associated with higher odds of vaccination delay and lower odds of full coverage (Dombkowski, Lantz and Freed 2004). In North-Eastern Libya, a woman's parity has been positively associated with the number of children deceased (Bhuyan 2000), and among women 15 – 24 years in India, marriage before the age of 18 years has been associated with higher adjusted odds of stunting and underweight in the next generation (Raj et al. 2010). In 6 of 15 lower-income countries, infants born to adolescent mothers have had lower odds of immunization than infants born to older mothers (Reynolds et al. 2006). Studies also have documented an inverse association between measures of fertility and children's schooling, suggesting that children in smaller families with less (sibling) competition for resources are more likely to have better access (Adamchak and Ntseane 1992; de la Croix and Doepke 2003), attainment (Booth and Kee 2009), and performance (e.g., Hanushek 1992).

LOWER FERTILITY, LATER FIRST CHILDBEARING, AND GENDER GAPS IN WELL-BEING

The *relative* benefits accrued to girls versus to boys from the "pro-quality norms" that emerge with declining fertility are likely to depend on prevailing gender norms (e.g., Das Gupta and Bhat 1997; Das Gupta and Shuzhuo 1999). In their study of the effects of fertility decline on gender bias in India, Das Gupta and Bhat (1997) described two potentially competing effects on girls' excess mortality in populations with strong preferences for sons. We extend their arguments to a general discussion of how declines in fertility may influence gender gaps in well-being in these and other contexts.

In populations with parents who prefer an equal investment strategy, sons and daughters should benefit similarly from smaller family sizes, and in the aggregate, any pre-existing gender gap in well-being should remain stable (unless parents initially compensate for past inequities). In populations having stronger preferences for sons, smaller family sizes may affect gender gaps in well-being through two channels. First, if girls are disadvantaged relative to boys at higher parities, then smaller average family sizes should improve girls' well-being, ceteris paribus. Das Gupta and Bhat (1997) have called this pattern a *parity effect*, and we call it here a *gendered parity effect* to distinguish it from the general improvements to children's well-being that may arise from reductions in family size. This typology stems from South Asian data showing that girls' excess mortality has concentrated in the higher parities (e.g., Das Gupta 1987; Muhuri and Preston 1991; Pebley and Amin 1991).

Although the gendered parity effect is a purely structural effect of falling fertility, the actual distribution of newly available resources will likely depend on cultural norms. In populations having stronger preferences for sons, an *intensification effect* occurs when boys accrue a larger share of these resources. Such effects are based on observations in some countries, where girls' excess mortality at a given parity has amplified as total fertility has fallen (Hull 1990; Zeng Yi et al. 1993). One explanation for this effect is that norms favoring a smaller family size change more rapidly than do norms about the preferred number of sons (Das Gupta and Bhat 1997). The differential pace of these normative changes results in greater "pressure" at each parity to invest in sons. Das Gupta and Bhat (1997) have provided evidence for India that the pace of fertility decline has exceeded declines in women's desired number of sons.[8]

As fertility declines in populations having stronger son preference, the gendered-parity and intensification effects are counter-balancing forces, and their net effect on gender gaps in well-being is an empirical question. Scenarios (a) – (d) in Figure 2 depict four possible trajectories of the net effects of these competing forces as well as likely scenarios in populations lacking strong son preference but still showing some gender gaps in well-being. In all scenarios, we assume for the sake of argument that some gender gap in well-being exists at the outset, with boys having an initial advantage over girls.

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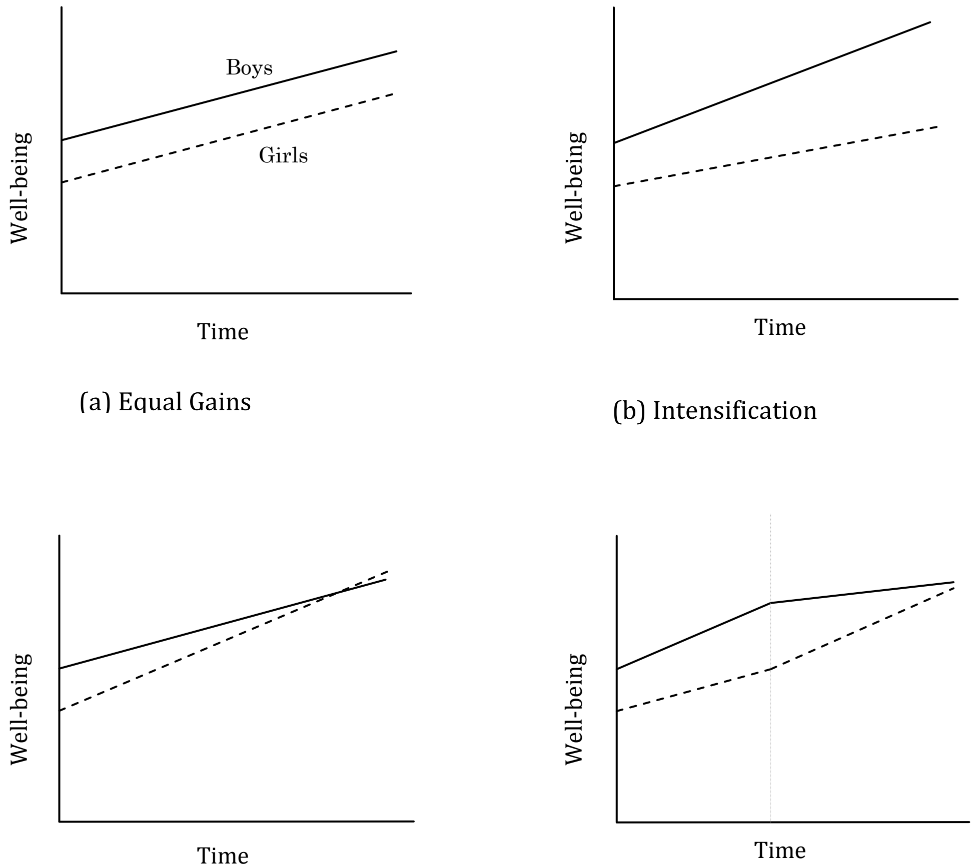
[8] According to the authors (p. 307), "in India during the 1980s, … Total Fertility fell by 20 per cent (Registrar General of India 1981 and 1991), whilst the number of sons desired by women who did not have any fell by only 7.4 per cent (Operations Research Group 1980 and 1988)."

Figure 2: Four Scenarios Depicting Likely Trajectories of Gender Gaps in Well-Being Following Declines in Fertility



(a) Equal Gains

(b) Intensification

(c) Gendered parity effect, changing gender norms

(d) Intensification followed by changing gender norms

NOTE: Adapted from Das Gupta and Bhat (1997)

Scenario (a) depicts the constant gender gap in well-being that results from boys and girls accruing the benefits of fertility decline at similar rates. In populations with stronger preferences for sons, this plot would be evidence of an equal influence of the gendered-parity and intensification effects. In populations with more balanced gender preferences, this plot could be evidence of a biological advantage of boys. Scenario (b) depicts a predominant intensification effect, in which boys accrue the benefits of declining fertility more rapidly than do girls. In their study of child mortality in India in the 1980s, Das Gupta and Bhat (1997) attributed an increasing number of excess deaths to girls as evidence of a predominant intensification effect. Because son (or male) preference is more common than daughter (or female) preference in poorer countries (Fuse 2010; World Values Survey 2011) and family-size norms may change more rapidly than gender norms, an intensification of gender gaps in well-being *may* occur even where initial levels of son preference are not extreme.

Scenario (c) depicts the situation in which girls accrue the benefits of declining fertility more rapidly than do boys. In populations with stronger preferences for sons, this effect could arise because the gendered parity effect outweighs the intensification effect. More likely, this scenario arises because declining fertility coincides with shifting gender norms in which women and girls become more equally valued (McDonald 2000). The crossover of the gender gap in well-being in Scenario (c) shows how changing norms could lead to a long-run advantage for women or girls at least for some aspects of well-being. The increasingly widespread emergence of higher schooling attainment for ever-enrolled girls than boys exemplifies this crossover (Knodel 1997; Grant and Behrman 2010). Finally, Scenario (d) depicts changing gender gaps in well-being that may arise in societies where son preference is strong initially but diminishes. In this scenario, the intensification effect dominates at first, resulting in a widening of the gender gap in well-being. Later, this gap narrows as changes in gender norms result in faster improvements for girls than boys. Trends in the sex ratio at birth in South Korea have followed this trajectory, although Chung and Das Gupta (2007) have argued that urbanization and industrialization (rather than declining fertility) were the reasons that the population's preference for sons "unraveled".

Three expectations follow from the foregoing discussion. First, declining fertility and especially increasing ages at first birth should be associated with better aggregate maternal nutrition, a universal need for women's well-being. Second, an environment of declining fertility and later first childbearing should encourage families to invest more in each pregnant mother and child, which would be reflected in higher levels of women's and girls' met needs for well-being (e.g., prenatal care, trained attendance at delivery, vaccination coverage, and schooling) as well as reductions in the objective harms associated with unmet needs (e.g., mortality and malnutrition). Third, although girls' well-being should improve with declining fertility and later first childbearing, the *relative* benefits accrued to girls versus boys will depend on prevailing gender norms. A preference for equal investments would result in stable gender gaps; a predominant gendered-parity effect would result in diminished gender gaps; a dominant intensification effect would result in rising gender gaps; and an initial intensification effect followed by changing gender norms would result in rising then diminishing gender gaps.

SAMPLE AND DATA

For this analysis, the country was the unit of observation, and the main sources of data were the Demographic and Health Surveys (DHS) (Measure DHS 2011) and the World Bank's World Development Indicators (WDI) database (World Bank 2011). The sample sizes for our analyses included 60 – 75 countries in which 124 – 187 DHS had been conducted from 1985 to 2008. The average number of observations per country ranged from 2.0 to 2.5 across analyses.

Our outcomes were derived from the DHS and included three sets of variables that captured harms associated with unmet needs for well-being (e.g., mortality, malnutrition) and benefits associated with met needs for well-being (e.g., nutrition, access to health care, schooling) for (a) women, (b) girls, and (c) girls relative to boys. For *women's well-being*, one measure of malnutrition captured the percentage of non-pregnant but recently delivered 15 – 49 year-olds who were *underweight*, or with a body-mass index (BMI) in kilograms per meters squared below 18.5. This measure included women who had had a birth in the prior three years and were more than three months postpartum.[9] Two measures of women's access to health care captured the percentages of births in the three years before each survey for whom the mother had: any prenatal care and any professional attendant at delivery. Professional attendants included only medically trained and licensed personnel, such as doctors, nurses, trained midwives, and trained auxiliary health personnel.

For *girls' well-being*, our measure of mortality captured the number of deaths to girls 1 – 4 years per 1,000 girls 1 – 4 years in the 10 years before the DHS survey.[10] Four measures of (mal-) nutrition for girls less than 36 months at the time of each DHS captured the mean height-for-age (*haz*) and weight-for-age (*waz*) *z*-scores with respect to the WHO/CDC/NCHS international reference population (Rutstein and Rojas 2006), the percentage *stunted* or below minus two standard deviations (SD) from the median *haz* in the reference, and the percentage *underweight* or below minus two SD from the median *waz* score in the reference. The measure of girls' access to health

[9] We considered a measure for the percentage of pregnant and non-pregnant women with any (mild, moderate, or severe) anemia; however, only 15 countries had data for at least two time points and only 35 countries had data for at least one time point.

[10] A measure for early-child mortality averaged over a shorter interval of time was not available from STAT-COMPILER, the online national statistics database for the DHS.

care captured the percentage of those 12 – 23 months who had received all recommended vaccines by the time of the survey (according to the child's vaccination card or mother's report). These vaccines included BCG, DPT 1 – 3, Polio 0 – 3, and Measles. Finally, the measure of schooling captured the percentage of girls 11 – 15 years in the household population who were attending school at the time of the survey.

Measures of *gender gaps in well-being* captured the difference between girls and boys in their: risks of mortality at ages 1 – 4 years, *haz* and *waz* scores at 0 – 36 months, percentages stunted and wasted at 0 – 36 months, percentages fully vaccinated at 12 – 23 months; and percentages attending school at 11 – 15 years. For each gap, the difference between comparable measures for boys and girls was taken so that a positive value reflected a disadvantage for girls (e.g., positive values for *girls'* minus *boys'* risks of mortality and for *boys'* minus *girls' haz* scores reflected disadvantages for girls).

The *explanatory variables* also were derived from the DHS and included two aggregate measures of the fertility regime. The total fertility rate (TFR), or the total number of births per woman of reproductive age, was estimated from age-specific fertility rates for women 15 – 49 years in the period 0 – 4 years before each DHS. The median age at first birth was estimated for women 25 – 49 years at the time of each DHS.[11] To permit non-linear associations between changes in fertility and changes in well-being, we included quadratic terms for our measures of fertility in all models. However, in only 3 of the 17 adjusted models with women's median age at first birth was its quadratic term significant, and in only 9 of the 17 adjusted models with total fertility was its quadratic term significant. For these nine outcomes, an inverse relationship between total fertility and well-being was observed across the two model specifications, but models with the quadratic term suggested that the gains in well-being were greater at lower levels of fertility. Overall, given the comparability of the findings for a large majority of the outcomes, the more parsimonious models excluding the quadratic terms are presented (and models with the quadratic terms are available upon request). Finally, we also considered as explanatory variables age-specific fertility rates to capture declines in fertility resulting from restriction at older ages (e.g., 40 – 44 years) or from delayed childbearing or restriction at younger ages (e.g., 20 – 24 years). The results from these analyses (available upon request) suggested that reductions in fertility at younger ages were

[11] We explored models using the median age at first birth for women 30 – 34 years of age at the time of each DHS, and the results were generally comparable.

associated most strongly with improvements in girls' well-being and that gender gaps in children's well-being responded most to reductions in fertility at older ages. In general, the results were analogous to those using TFR, and so those for TFR and the median age at first birth are presented.

The main *control variables* included country- and time-fixed effects, which captured, respectively, unobserved time-*invariant* national attributes (such as language) and unobserved time-*variant* national attributes (such as political regime) as captured by the years of each DHS for a given country. Four other time-varying national attributes were included to capture socioeconomic changes that may have been correlated with changes in aggregate fertility and well-being (see Chung and Das Gupta 2007). Data for these variables were obtained from the World Development Indicators database (World Bank 2011). One measure of technological innovation captured mobile cellular subscriptions per 100 people in the same year as each DHS.[12] One measure of urbanization captured the percentage of the total population that was living in urban areas, as defined by the national statistical offices and estimated for the same year as or up to six years before each DHS.[13] Two measures of national economic conditions captured five-year averages of the Gross Domestic Product (GDP) per capita and the net Official Development Assistance (ODA) per capita received, both in current U.S. dollars.[14] For these two measures, annual figures for the 0 to 4 years preceding each DHS were averaged to match the years for which TFRs were calculated.[15]

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[12] Mobile cellular subscriptions were defined as subscriptions to a public mobile telephone service using cellular technology that provides access to the public switched telephone network. Postpaid and prepaid subscriptions were included, and WDI sources for subscription data were taken from the International Telecommunication Union, World Telecommunication/ICT Development Report and database, and World Bank estimates.

[13] Values were derived from World Bank population estimates and urban ratios from the United Nations World Urbanization Prospects.

[14] For the average GDP per capita, WDI sources included World Bank national accounts data and OECD national accounts data files. The WDI sources for the net ODA were taken from the Development Assistance Committee (DAC) of the Organization for Economic Co-operation and Development, Geographical Distribution of Financial Flows to Developing Countries, Development Co-operation Report, and the International Development Statistics database. According to the World Bank (2011), the Net ODA includes disbursements of loans made on concessional terms (net of repayments of principal) and grants by official agencies of the members of the DAC, by multilateral institutions, and by non-DAC countries to promote economic development and welfare in countries and territories in the DAC list of ODA recipients. It includes loans with a grant element of at least 25 percent (calculated at a rate of discount of 10 percent).

[15] We also considered controlling for within-country changes over time in the criteria for including women in the DHS (e.g., ever-married only versus all women 15 – 49 years); however, these criteria remained

METHODS

First, we examined univariate distributions of all outcomes, explanatory variables, and covariates to assess their completeness and distributions. We then explored within-country trends in all outcomes, explanatory variables, and covariates to ensure sufficient change over time to permit time-series analyses. Finally, we examined bivariate plots of all outcomes and explanatory variables to explore potential non-linearities in their associations.

Preliminary multivariate models for each outcome were estimated using linear panel regression with country- and time-fixed effects. In general, fixed-effects models explore the relation between an explanatory variable and an outcome within an entity, in this case a country. Each country is assumed to have unique, time-invariant attributes that may influence the explanatory and outcome variables, confounding their estimated relationship. Fixed-effects models remove these effects of unobserved time-invariant attributes so that the "net influence" of the explanatory variable can be assessed. The general equation for the country fixed-effects model is:

$$Y_{it}= \beta_1 F_{it} + \alpha_i + u_{it} \qquad (1)$$

where Y_{it} is the dependent variable (such as the percentage of girls 11 – 15 years attending school) in country i at time t, F_{it} denotes the main explanatory variable (such as the total fertility rate) for country i at time t, β_1 denotes the coefficient for that explanatory variable, α_i (I = 1….n) is the unknown intercept for each country (n country-specific intercepts), and u_{it} is the error term. To control for unexpected time-variant effects or special events that may affect the outcome, we then estimated three other models that adjusted the base model reflected in equation (1) for: (2) time fixed-effects, (3) time-variant national socioeconomic conditions, and (4) time fixed-effects and time-variant national socioeconomic conditions, as follows:

$$Y_{it}= \beta_1 F_{it} + \beta_T T_t + \alpha_i + u_{it} \qquad (2)$$

$$Y_{it}= \beta_1 F_{it} + \beta_E E_{it} + \alpha_i + u_{it} \qquad (3)$$

$$Y_{it}= \beta_1 F_{it} + \beta_T T_t + \beta_E E_{it} + \alpha_i + u_{it} \qquad (4)$$

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consistent across survey years for all countries except India, Morocco, Nepal, and Turkey, in which never married women were added in later years. So, the country-fixed effects controlled sufficiently for variation in the sample across countries.

where T_t is time as a design variable and t-1 time periods entered into the model and E_{it} denotes a vector of national socioeconomic conditions for country i at time t. Robust standard errors were estimated to account for clustering at the county-level, and the population sizes for each country averaged over the period of analysis were used as analytic weights (Dorius 2008).

Notably, the fixed-effects model assumes that time-invariant attributes are unique to the entity under analysis (e.g., a given country), so the country's error term and the constant capturing the country's attributes should not be correlated with those for other countries. If the error terms across countries are correlated, then fixed-effects modeling is not suitable, and the relationship may need to be modeled using random effects. To explore this possibility, we estimated random-effects models following the sequential model-fitting strategy depicted in equations (1) – (4) (in this case, α_i became the random effect) and again using population-average analytic weights.[16] We estimated the relative fits of fixed- and random-effects models by performing a Hausman test on models that did not include average population sizes as analytic weights (Hausman 1978). A rejection of the null hypothesis would lead one to conclude that the fixed-effects models were preferable to the random-effects models. In this analysis, only 4 of the 34 Hausman tests were significant at $p \leq 0.05$, and in only one of these four cases did the inference for total fertility differ; therefore, the random-effects estimates are presented in the main results, and the fixed-effects estimates are presented in appendices and discussed where appropriate.

As a last step in the main analysis, we re-ran all final random-effects models excluding the observations for India. This step had several rationales. First, India was more heavily weighted in the full-sample analysis, being five times more populous than the next most populous sample country (Indonesia, with an average population size of 195 million for the period). Second, this exclusion enabled us to examine whether the estimated relationships of fertility decline and later childbearing with changes in well-being varied when a high-son-preference setting was included or excluded. (We recognize that India is a diverse country with considerable variation in levels of son preference [e.g., Dyson and Moore 1983]; however, compared to 50 other countries with DHS data, India ranks 11th in terms of the percentage of women who have reported a preference for sons [Fuse 2010]).

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[16] Robust standard errors, however, were not estimable with the random effects model that included population-average weights. For unweighted random effects models, we compared the results of a model estimated without robust standard errors to models accounting for clustering within countries and regions. The inferences were comparable across these three models.

Thus, removing a more heavily weighted country with high reported son preference permitted us to observe whether there were meaningful changes in our inferences regarding the effects of fertility decline on women's and girls' well-being and gender gaps in children's well-being. A priori, we expected that removing the case of India would: (1) strengthen any observed benefits of fertility decline for women's and girls' well-being and (2) weaken any observed intensification effects of fertility decline on gender gaps in children's well-being.

Finally, several diagnostics for model fit and robustness were performed. First, F-tests for the joint significance of the design variables for time indicated that their inclusion was warranted. Second, variance inflation factors (VIFs) estimated from ordinary least squares regressions controlling for economic indicators and time fixed effects were well below the standard cutoff of 10 for the total fertility rate (1.8 to 2.4), the median age at first birth (2.0 to 2.7), and the four economic indicators (1.2 to 6.6), suggesting that multi-collinearity was not influencing the least square estimates (Hair et al. 1995; Marquardt 1970; Neter, Wasserman and Kutner 1989; O'Brien 2007). Third, we explored variation in the relationship between F_{it} and Y_{it} in equation (4) by interacting measures of fertility with three sets of variables (full results available upon request). Interactions of total fertility and the median age at first birth with a regional indicator for Sub-Saharan Africa (1 = yes, 0 = no) (Kögel 2004)[17] were significant for only a subset of our 17 outcomes (6 and 11, respectively), and in these cases, the benefits of fertility decline and later childbearing tended to be weaker in Sub-Saharan Africa than elsewhere. Interactions with an indicator for any son preference at baseline (defined as a sex ratio at birth > 1.05 versus ≤ 1.05 at the time of the first DHS)[18] were not significant in 14 of 17 models with the total fertility rate and in 16 of 17 models with the median age at first birth. In interactions with our socio-economic variables, the benefits of declining fertility on some aspects of girls' and women's well-being were boosted in more urban populations with broader cell phone use (e.g., Chung and Das Gupta 2007; Das Gupta and Shuzhuo 1999), and the benefits of later first childbearing for girls relative to boys tended to be less pronounced with increasing net ODA per capita (results available upon request). Finally, our main analysis was restricted to the countries in which the DHS have been conducted. In some cases, countries that were surveyed in earlier years were not surveyed in later years. If a country's observation (or not) in the sample was associated with changes in its population's aggregate fertility and well-being, then the relationships estimated here

[17] This demarcation also arguably captures different regional patterns and timings of fertility change.

[18] 18% of the observations in this analysis had initial sex ratios greater than 1.05.

may be biased. We assessed the effects of country attrition on our inferences by imputing observations for selected outcomes using data from the Multiple Indicator Cluster Surveys (MICS) (UNICEF 2011). Re-estimation of models with imputed data for some countries showed no meaningful differences in our inferences.

TABLE 1. Descriptive Statistics for Measures of Women's and Girls' Well-being, Gender Gaps in Children's Well-being, and Explanatory and Contextual Variables, Demographic and Health Surveys (DHS) 1985-2008

	1990-1994					2005-2008					Total						
	Mean	SD	Min	Max	N[a]	Mean	SD	Min	Max	N[a]	Mean	SD	Min	Max	N[a]	n[b]	N/n[c]
Women's Well-being																	
% with any prenatal care, births in prior 3 y	65.2	17.2	27.6	97.2	30	76.9	14.3	27.8	99.5	39	71.5	17.8	27.0	99.8	177	72	2.5
% with trained attendant at delivery, births in prior 3 y[d]	39.4	16.9	9.5	93.1	30	53.2	18.9	19.0	99.5	36	47.7	20.3	8.6	99.6	175	73	2.4
% with BMI < 18.5, non-pregnant women with birth in prior 3 y[e]	6.1	5.2	1.3	19.4	17	28.9	14.7	0.7	39.9	32	25.6	16.5	0.6	52.0	124	60	2.1
Girls' Well-being																	
Deaths to girls 1–4 y per 1,000 1–4 y	43.5	24.6	5.6	231.8	30	29.2	23.0	1.0	135.5	39	37.1	27.0	1.0	231.8	186	75	2.5
% 12–23 m with specified vaccines[f]	43.6	13.8	14.5	86.2	30	51.0	19.1	18.2	92.4	36	47.8	17.8	1.1	92.4	176	73	2.4
Mean *haz* score < 3 y[g]	-1.6	0.4	-1.9	-0.6	25	-1.3	0.3	-1.7	-0.1	32	-1.4	0.4	-2.3	-0.1	152	67	2.3
% < 3 y stunted[h]	40.3	9.8	14.0	47.0	25	34.0	8.3	6.6	42.5	32	36.1	11.1	4.5	56.7	152	67	2.0
Mean *waz* score < 3 y[i]	-1.5	0.5	-1.9	-0.2	25	-1.4	0.6	-1.8	0.0	32	-1.4	0.6	-2.0	0.2	152	67	2.0
% < 3 y underweight[j]	40.3	15.0	4.7	50.9	25	37.1	14.7	3.9	47.0	32	35.2	16.2	2.1	52.7	152	67	2.3
% 11-15 y attending school	57.5	11.5	12.8	90.8	28	79.8	14.8	26.9	95.6	14	66.8	15.5	12.8	99.1	125	60	2.1
Gender Gaps in Children's Well-being (a positive value reflects girls' disadvantage)																	
Gap (girls–boys 1–4 y) in risk of mortality	6.2	9.3	-16.3	20.2	30	4.3	5.4	-13.8	9.2	39	4.3	6.9	-24.7	20.2	186	75	2.0
Gap (boys–girls 12–23 m) in % with specified vaccines	2.0	2.6	-10.1	7.8	30	1.9	3.1	-6.3	11.5	36	1.3	3.2	-12.0	28.5	176	73	2.4
Gap (boys–girls < 3 y) in mean *haz* scores[g]	-0.1	0.0	-0.2	0.0	25	0.0	0.1	-0.3	0.3	32	0.0	0.1	-0.5	0.3	152	67	2.3
Gap (girls–boys < 3 y) in % stunted[h]	-1.4	1.2	-6.9	1.1	25	-0.4	2.5	-7.4	3.1	32	-0.8	2.7	-7.4	3.6	152	67	2.3
Gap (boys–girls < 3 y) in mean *waz* scores[i]	-0.1	0.1	-0.2	0.0	25	0.0	0.1	-0.3	0.2	32	0.0	0.1	-0.4	0.2	152	67	2.3
Gap (girls–boys < 3 y) in % underweight[j]	-1.9	1.0	-5.3	2.5	25	0.7	2.4	-8.6	4.3	32	-0.2	2.6	-8.6	4.3	152	67	2.0
Gap (men–women 11-15 y) in school attendance	14.4	10.0	-9.8	23.7	28	-1.9	4.8	-6.2	14.0	14	8.4	9.6	-9.8	45.2	125	60	2.1
National Fertility Regime																	
Total Fertility Rate, Women 15–49 y 0–4 y before survey	3.9	1.0	2.6	7.3	30	3.4	1.2	1.6	7.1	38	3.7	1.2	1.6	7.5	184	74	2.5
Median Age at First Birth, Women 25 – 49 y	20.0	1.2	17.5	22.8	31	20.3	1.3	17.9	23.9	39	20.2	1.4	17.2	24.0	186	75	2.5
National Economic Conditions																	
Mobile cellular subscriptions per 100 population, same year as DHS	0.0	0.0	0.0	0.2	31	21.1	19.0	0.5	99.9	39	6.2	13.2	0.0	99.9	187	75	2.0
Avg GDP per capita 0 - 4 y prior, current US$[k]	653.6	661.7	188.9	2613.5	31	781.5	493.1	131.4	3227.7	39	789.9	760.7	131.4	5566.0	187	75	2.5
Urban population as a % of total, current or prior year to DHS	33.8	15.3	5.4	74.8	31	34.4	12.3	12.5	78.3	39	35.7	16.0	5.2	80.1	187	75	2.0
Avg net ODA received per capita 0-4 y prior, current US$[l]	11.7	18.1	1.4	183.7	31	12.3	18.1	1.3	135.5	39	12.7	17.6	0.9	183.7	187	75	2.5

NOTE.

[a] N refers to number of DHS.

[b] n refers to number of countries.

[c] N/n refers to the number of DHS per country.

[d] Trained attendants included only medically trained and licensed personnel, such as doctors, nurses, trained midwives, and trained auxiliary health personnel.

[e] BMI is measured in units kg/m^2.

[f] Specific vaccinations include: BCG, DPT 1 – 3, Polio 0 – 3, and Measles.

[g] *haz* refers to height-for-age z-score.

[h] % stunted refers to the percentage below minus two standard deviations (SD) from the median *haz* in the WHO/CDC/NCHS international reference population.

[i] *waz* refers to weight-for-age z-score.

[j] % underweight refers to the percentage below minus two standard deviations (SD) from the median *waz* score in the WHO/CDC/NCHS international reference population.

[k] GDP refers to Gross Domestic Product.

[l] ODA refers to Overseas Development Assistance.

The top panel of Table 1 shows descriptive statistics for measures of women's and girls' well-being and gender gaps in children's well-being. With respect to women's well-being, estimates for 1990 – 4 and for 2005 – 8 suggest a rise in the use of prenatal care and in having a trained attendant at delivery between the two periods. In 1990 – 4, on average, women had some prenatal care for 65% of births in the prior three years, and this average was 77% in 2005 – 8. The mean prevalence of having had a trained attendant at delivery for births in the prior three years was 39% in 1990 – 4 and 53% in 2005 – 8. The average percentage of non-pregnant, recently delivered women who were underweight was 6% in 1990 – 4 and 29% in 2005 – 8; however, this unexpected trend was not observed when the sample was restricted to countries with surveys in both periods).[19]

With respect to girls' well-being and gender gaps in children's well-being, girls' risk of dying in early childhood averaged 44 deaths per 1,000 girls in 1990 – 4 and 30 deaths per 1,000 girls in 2005 – 8. The mean risk of dying in early childhood was greater for girls than for boys in both periods, by 6 deaths per 1,000 in 1990 – 4 but by 4 deaths per 1,000 in 2005 – 8. The average percentage of girls 12 – 23 months receiving all recommended vaccinations rose from 44% in 1990 – 4 to 51% in 2005 – 8, but the gender gap in vaccination coverage consistently favored boys by about 2.0% in both periods. High mean percentages of girls 0 – 3 years were stunted (40%) and underweight (40%) in 1990 – 4; however, mean levels of malnutrition were lower by 1% – 2% for girls than for boys in this period. In 2005 – 8, the mean percentages of girls who were stunted and underweight were lower (34%, 37%) than in 1990 – 4, but they still exceeded one third, and gender gaps in stunting and underweight had either disappeared or reversed to reflect a slight disadvantage for girls (see positive gender gap in underweight in 2005 – 8). On average, 58% of girls were attending school in 1990 – 4, and the mean gender gap in this percentage was 14% in favor of boys. The surveys conducted in 2005 – 8 suggest a reversal in the gender gap in schooling, with an average of 80% of girls 11 – 15 years attending school and a mean gender gap of 2% favoring girls.

Comparing the average TFR for surveys in 1990 – 4 and 2005 – 8 suggests that fertility was declining at least for some countries between these periods. In 1990 – 4, the average TFR was 3.9, and the lowest TFR was above replacement (2.6 in Turkey in 1993). In contrast, among surveys in

[19] Bangladesh, India, and Eritrea, the three countries in our sample with the highest percentages of women with BMI < 18.5, were represented in the sample of countries for the 2005 – 8 period but were not represented in the 1990 – 4 period. Data on this variable were not collected in either the 1992 India DHS or the 1993 Bangladesh DHS.

2005 – 8, the average TFR was 3.4, and the lowest TFR was below replacement (1.6 in Moldova in 2005). The average median age at first birth among women 25 – 49 years was 20.0 years in 1990 – 4 and 20.3 years in 2005 – 8.

According to the average GDP per capita, national economic conditions improved between the periods, with GDP per capita averaging $654 in 1990 – 4 and $782 in 2005 – 8. The average percentage of the population living in an urban area and the average net ODA per capita did not change much between the periods, with the former at 34% and the latter at $12 in 1990 – 4 and 2005 – 8. Comparing the average number of mobile cellular subscriptions per 100 people across periods suggests rapid growth in this technology, from 0 in 1990 – 4 to 21 in 2005 – 8.

Figure 3: National Trends in Women's Well-Being and Gender Gaps in Well-Being (1985-2008)

FIGURE 3a. Trends in Total Fertility Rate, Women 15–49 y (0–4 y before survey) and Median Age at First Birth, Women 25 – 49 y, Demographic and Health Surveys for 1985-2008

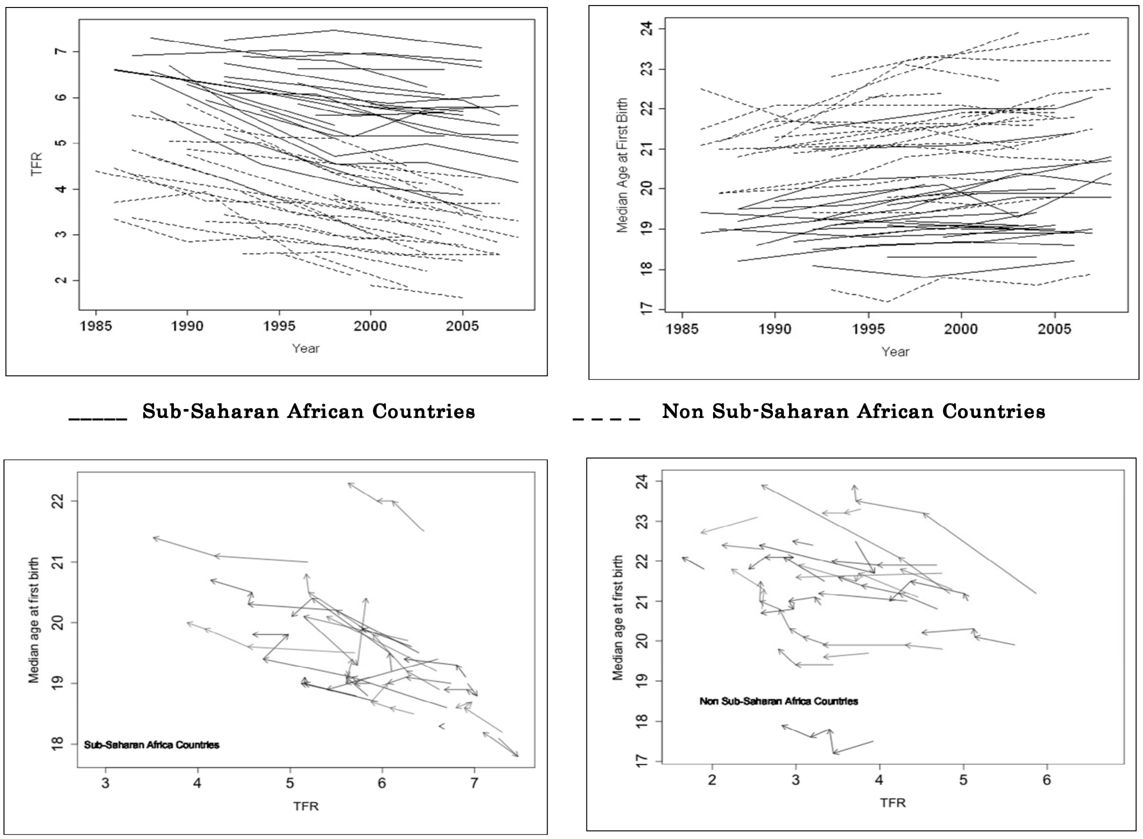
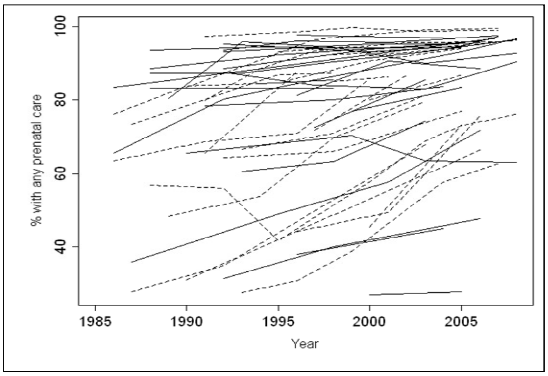
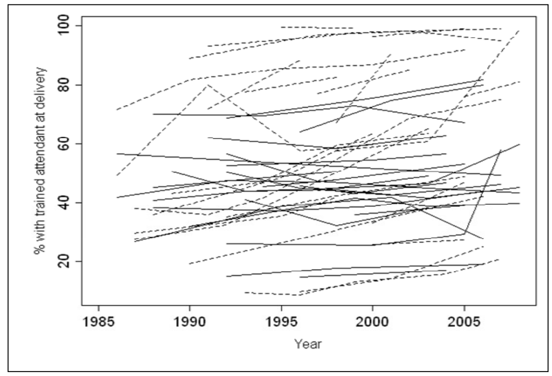
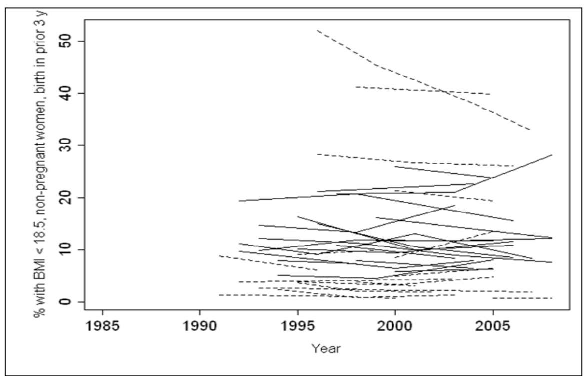


FIGURE 3b. Trends in Selected Measures of Women's Well-being, Demographic and Health Surveys for 1985-2008





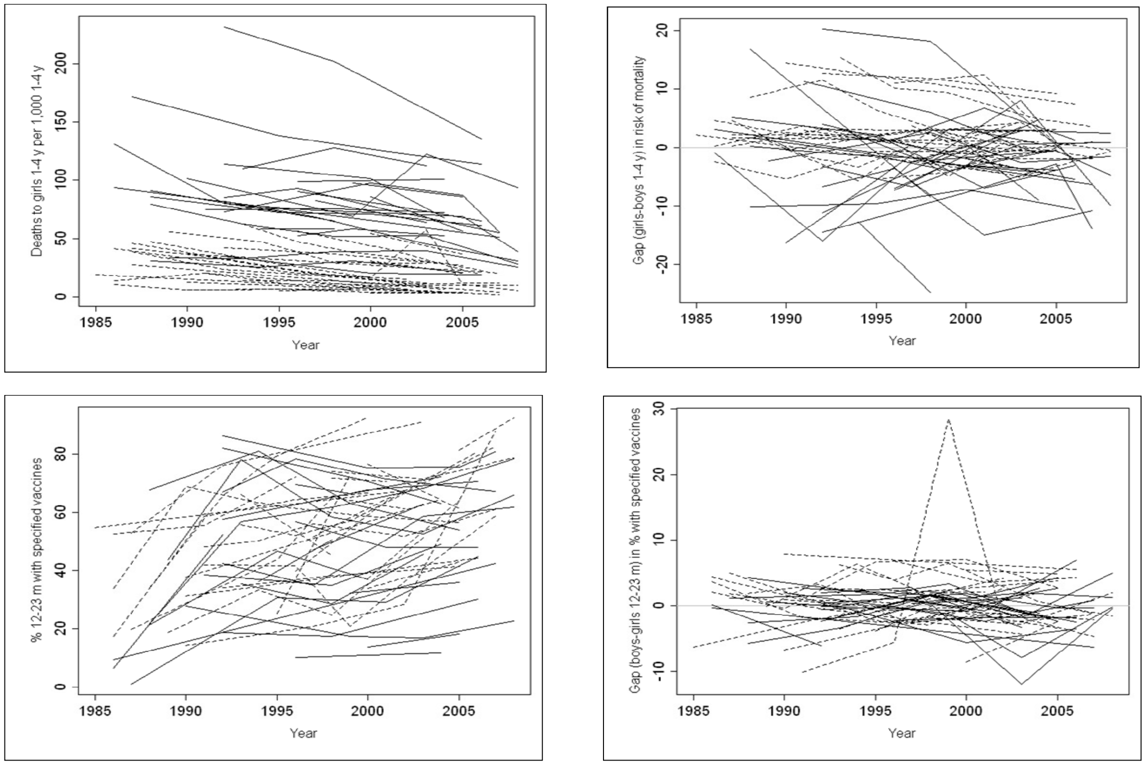
NOTE. Trained attendants included only medically trained and licensed personnel, such as doctors, nurses, trained midwives, and trained auxiliary health personnel.



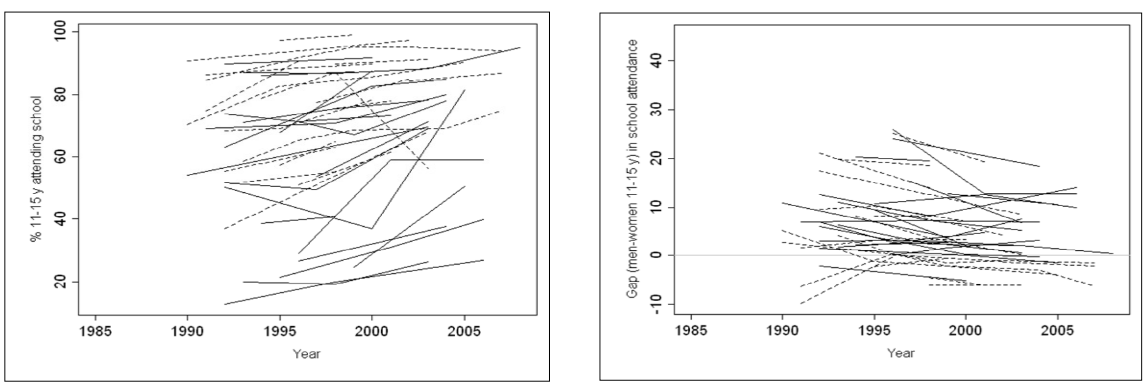
NOTE. BMI is measured in units kg/m^2.

_____ Sub-Saharan African Countries **_ _ _ _ Non Sub-Saharan African Countries**

FIGURE 3c. Trends in Selected Measures of Girls' Well-being and Gender Gaps in Well-being, Demographic and Health Surveys for 1985-2008.

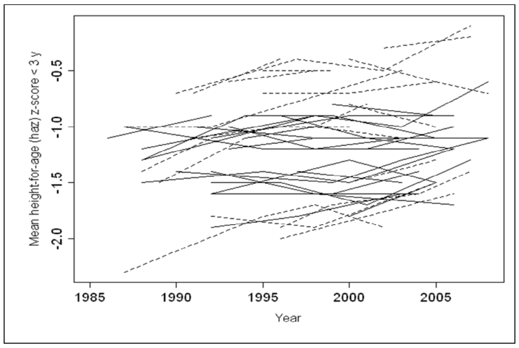


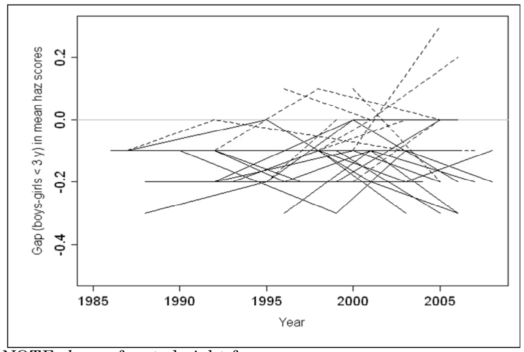
NOTE. Specific vaccinations include: BCG, DPT 1 – 3, Polio 0 – 3, and Measles.



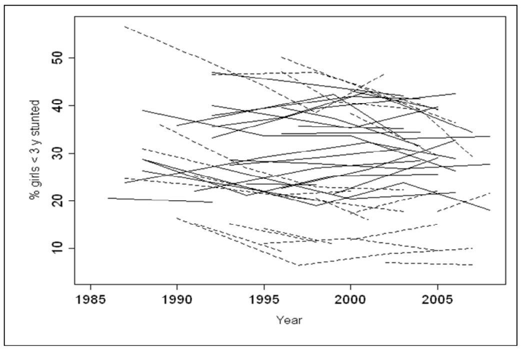
_____ Sub-Saharan African Countries _ _ _ _ Non Sub-Saharan African Countries

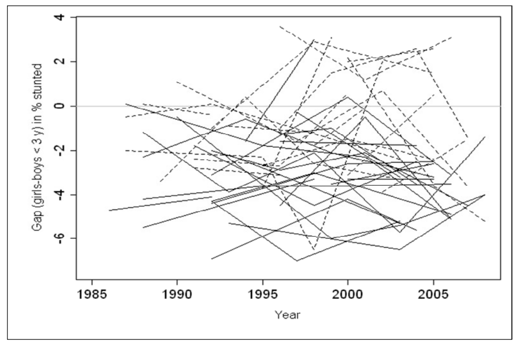
FIGURE 3c (continued)



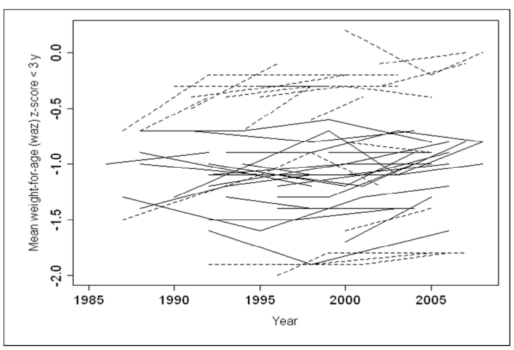


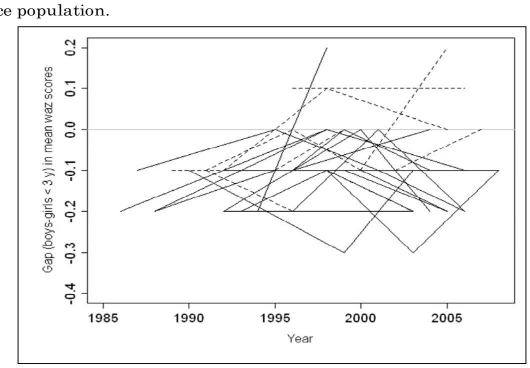
NOTE. *haz* refers to height-for-age z-score.



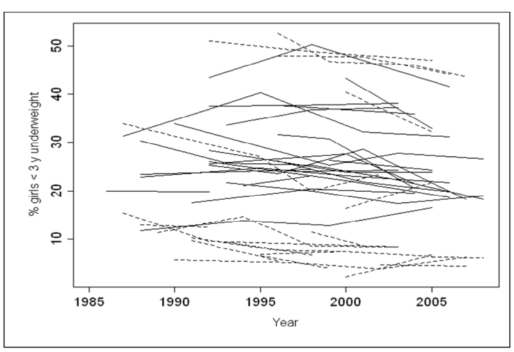


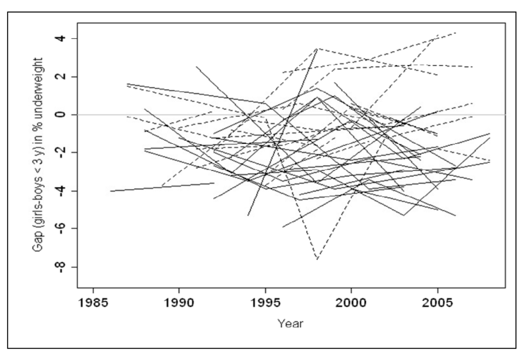
NOTE. % stunted refers to the percentage below minus two standard deviations (SD) from the median haz in the WHO/CDC/NCHS international reference population.





NOTE. *waz* refers to weight-for-age z-score.





NOTE. % underweight refers to the percentage below minus two standard deviations (SD) from the median waz score in the WHO/CDC/NCHS international reference population.

_____ Sub-Saharan African Countries **_ _ _ _ Non Sub-Saharan African Countries**

Figures 3a – 3c depict national trends in: measures of fertility and the relationship between these trends (3a), our three measures of women's well-being (3b), and our measures of girls' well-being and gender gaps in children's well-being (3c). In all graphs, the solid lines depict trends for countries in Sub-Saharan Africa, and the dotted lines depict trends for countries in other regions.

As expected (Figure 3a), levels of fertility are higher and median ages of first birth are lower in Sub-Saharan African countries than in other regions throughout the period (1985 – 2008). Still, most countries show declines in TFR over the period, although trends in the median age at first birth are more variable. On average, a slight increase in the median age at first birth is apparent across countries (see also Table 1), but trends in this variable are largely flat in those countries with the lowest initial median ages at first birth (e.g., Bangladesh, Chad, Mali, Uganda, and Zambia). As researchers have described, declines in total fertility in our sample typically have occurred earlier and faster than have increases in the median age at first birth.

With respect to trends in women's well-being (Figure 3b), the percentage of non-pregnant, recently delivered women who were underweight seems to have declined in some countries (most notably in Bangladesh); however, within-country reductions in this prevalence typically were modest. In most countries, women's access to maternal health care improved over the period, with increases in the percentages of births in the prior three years whose mothers received any prenatal care and delivered with a trained attendant.

With respect to national trends in girls' well-being (Figure 3c), marked improvements were broadly evident, as reflected in declining risks of 1 – 4 mortality, increasing vaccination coverage at 12 – 23 months, increasing average *haz* and *waz* scores at 0 – 36 months, corresponding reductions in the percentages stunted and underweight at this age, and increasing school attendance at 11 – 15 years. That said, national levels of and trends in gender gaps in these measures generally tell two distinctive stories. First, compared to girls in a majority of countries, boys initially fared better in terms of survival and access to services: they had lower risks of mortality at 1 – 4 years (in 41 of 75 countries), higher rates of vaccination coverage at 12 – 23 months (in 45 of 73 countries), and higher rates of school attendance at 11 – 15 years (in 47 of 60 countries). In most of the countries where these gender gaps initially favored boys, a narrowing of the gaps typically was apparent. Second, compared to boys in a majority of countries, girls initially fared slightly better in terms of nutritional status: compared to boys 0 – 36 months, girls initially had higher mean *haz* scores (in 48 of 67 countries), lower rates of stunting (in 58 of 67 countries), higher mean *waz* scores (in 45 of 67 countries), and lower rates of underweight (in 45 of 67 countries). In some of the countries where these gender gaps initially favored girls, a narrowing of the gaps typically was apparent, meaning that girls tended to lose their nutritional advantage over time (e.g., because improvements in nutrition were faster for boys). Of the 41 countries with longitudinal measures of child nutrition, girls initially had lower rates of stunting and underweight in a majority of countries (35 and 30, respectively). Comparing the first and last observations within countries, the gender gap in these measures either narrowed or reversed over time in 17 countries for both of these outcomes.

Random-Effects Models of Women's and Girls' Well-Being and Gender Gaps in Well-Being

Table 2 presents the results of the random-effects models exploring to what extent changes in total fertility and the median age at first birth were associated with changes in women's and girls' well-being. Table 3 presents a similar set of results for measures of gender gaps in well-being. In these two tables, Column (1) depicts the results of the first series of these random-effects models, which account only for country effects. Controls for time effects and socioeconomic changes are included in the models presented in Column (2) and Column (3) respectively. Column (4) presents the results for the series of models with all three sets of controls, and we focus our discussion on these results.

TABLE 2. Populated-Weighted Random-Effect Estimates of Total Fertility Rate (TFR) and Median Age at First Birth (AFB) on Women's and Girls' Well-being, Demographic and Health Surveys for 1985-2008

	(1) B (*se*) *p*	(2) B (*se*) *p*	(3) B (*se*) *p*	(4) B (*se*) *p*	R^2	N/n
Explanatory Variable: TFR, women 15–49 y 0–4 y before survey						
Women's Well-being						
% with any prenatal care, births in prior 3 y	-11.14 (1.18) ***	-7.74 (1.47) ***	-8.31 (1.35) ***	-6.45 (1.63) ***	0.08	175/71
% with trained attendant at delivery, births in prior 3 y	-12.16 (1.36) ***	-6.84 (1.72) ***	-6.46 (1.20) ***	-4.23 (1.60) **	0.54	175/73
% with BMI < 18.5, non-pregnant women with birth in prior 3 y	-2.05 (1.06) !	-3.01 (1.12) **	-4.86 (1.01) ***	-5.85 (0.97) ***	0.71	124/60
Girls' Well-being						
Deaths to girls 1–4 y per 1,000 1–4 y	18.94 (1.29) ***	17.21 (1.65) ***	17.77 (1.49) ***	17.59 (1.78) ***	0.69	184/74
% 12–23 m with specified vaccines	-10.35 (1.28) ***	-6.83 (1.54) ***	-9.92 (1.43) ***	-8.31 (1.71) ***	0.27	176/73
Mean *haz* score < 3 y	-0.14 (0.03) ***	-0.08 (0.04) *	-0.02 (0.03)	0.00 (0.03)	0.66	152/67
% < 3 y stunted	3.14 (0.79) ***	1.45 (0.90)	0.34 (0.68)	-0.40 (0.74)	0.73	152/67
Mean *waz* score < 3 y	-0.17 (0.03) ***	-0.08 (0.04) !	-0.03 (0.03)	0.03 (0.04)	0.54	152/67
% < 3 y underweight	3.82 (0.61) ***	0.62 (0.95)	0.49 (0.69)	-1.08 (0.91)	0.61	152/67
% 11-15 y attending school	-10.56 (1.04) ***	-4.58 (1.30) ***	-8.82 (1.15) ***	-3.83 (1.35) **	0.47	125/60
Explantory Variable: AFB, Women 25 – 49 y						
Women's Well-being						
% with any prenatal care, births in prior 3 y	9.63 (1.18) ***	5.50 (1.22) ***	6.18 (1.54) ***	5.18 (1.44) ***	0.20	177/72
% with trained attendant at delivery, births in prior 3 y	15.74 (1.16) ***	10.62 (1.29) ***	7.94 (1.31) ***	7.03 (1.31) ***	0.67	174/72
% with BMI < 18.5, non-pregnant women with birth in prior 3 y	-6.22 (0.89) ***	-5.93 (0.95) ***	-4.31 (1.14) ***	-3.83 (1.12) ***	0.64	124/60
Girls' Well-being						
Deaths to girls 1–4 y per 1,000 1–4 y	-15.65 (1.58) ***	-11.01 (1.74) ***	-11.70 (2.04) ***	-10.43 (1.95) ***	0.31	185/75
% 12–23 m with specified vaccines	7.39 (1.26) ***	5.27 (1.25) ***	5.89 (1.55) ***	5.80 (1.46) ***	0.24	175/73
Mean *haz* score < 3 y	0.26 (0.03) ***	0.20 (0.03) ***	0.12 (0.03) ***	0.12 (0.03) ***	0.70	152/67
% < 3 y stunted	-5.99 (0.71) ***	-4.37 (0.72) ***	-2.50 (0.71) ***	-2.09 (0.64) **	0.75	152/67
Mean *waz* score < 3 y	0.27 (0.03) ***	0.24 (0.03) ***	0.19 (0.03) ***	0.20 (0.03) ***	0.62	152/67
% < 3 y underweight	-5.78 (0.65) ***	-3.99 (0.74) ***	-3.81 (0.70) ***	-3.77 (0.72) ***	0.67	152/67
% 11-15 y attending school	7.15 (1.32) ***	4.74 (1.10) ***	2.53 (1.63)	3.51 (1.26) **	0.44	125/60
Controls (of country C, time T, economic change EΔ):	C	C, T	C, EΔ	C, T, EΔ		

NOTE.
Refer to the NOTE to Table 1 for detailed definitions of the variables.
! $p \leq 0.10$, * $p \leq 0.05$, ** $p \leq 0.01$, *** $p \leq 0.001$.

Women's well-being

As shown in Table 2, Column 4, a declining total fertility and an increasing median age at first birth generally were associated with improvements in women's well-being. Specifically, a decline of one birth in total fertility and an increase of one year in women's median age at first birth were associated with, respectively, increases of 6.5 and 5.2 percentage-points in the prevalence of prenatal care, as well as increases of 4.2 and 7.0 percentage points in the prevalence of trained attendants at delivery for births in the prior three years. Although a decline of one birth in total fertility was associated with an unexpected *increase* of 5.9 percentage points in the prevalence of underweight among non-pregnant, recently delivered women, a one-year increase in the median age at first birth was associated with an expected *decrease*–of 3.8 percentage points–in the prevalence of underweight in this group of women.

Girls' well-being

With respect to girls, declining total fertility and increasing median ages at first birth also was associated with substantial gains in well-being (Table 2, Column 4). Specifically, a decline of one birth in total fertility and an increase of one year in women's median age at first birth was associated with, respectively, 18 and 10 fewer deaths per 1,000 girls 1 – 4 years, increases of 8.3 and 5.8 percentage points in rates of full vaccination coverage at 12 – 23 months, and increases of 3.8 and 3.5 percentage points in rates of school attendance at 11 – 15 years. A decline in total fertility was not associated with changes in girls' nutrition, but a one-year increase in the median age at first birth was associated with increases of 0.12 and 0.20, respectively, in the mean *haz* and *waz* scores at 0 – 3 years as well as declines of 2.1 and 3.8 percentage points in rates of stunting and underweight at these ages.

Gender gaps in children's well-being

With respect to gender gaps in children's well-being (Table 3, Column 4), a decline in total fertility was associated with an *increase* in the gender gap in child mortality (which often initially favored boys) and a *narrowing* of the gender gaps in all measures of child nutrition (which often initially favored girls) (Table 5 summarizes these results visually). Compared to these results, those for models that included the median age at first birth seem contradictory: specifically, an increase in the median age at first birth was associated with a *decline* in the gender gap in early child mortality and a *widening* of the gender gap in measures of nutrition at 0 – 36 months. Recall that, in the models of girls' well-being, increases in the median age at first birth were associated with significant improvements in nutrition; however, these same associations were not significant when total fertility was the explanatory variable. We will return to these points in the Discussion.

TABLE 3. Populated-Weighted Random-Effect Estimates of Total Fertility Rate (TFR) and Median Age at First Birth (AFB) on Gender Gaps in Children's Well-being, Demographic and Health Surveys for 1985-2008

	(1)			(2)			(3)			(4)				
	B	(*se*)	*p*	B	(*se*)	*p*	B	(*se*)	*p*	B	(*se*)	*p*	R^2	N/n
Explanatory Variable: TFR, women 15–49 y 0–4 y before survey														
Gap (girls–boys 1–4 y) in risk of mortality	-0.757	(0.487)		-1.125	(0.520)	*	-1.162	(0.548)	*	-1.211	(0.603)	*	0.45	184/74
Gap (boys–girls 12–23 m) in % with specified vaccines	0.047	(0.258)		-0.032	(0.282)		-0.084	(0.302)		-0.229	(0.324)		0.24	176/73
Gap (boys–girls < 3 y) in mean *haz* scores	-0.035	(0.007)	***	-0.027	(0.009)	**	-0.050	(0.009)	***	-0.048	(0.011)	***	0.47	152/67
Gap (girls–boys < 3 y) in % stunted	-1.135	(0.165)	***	-0.968	(0.189)	***	-1.549	(0.190)	***	-1.615	(0.236)	***	0.62	152/67
Gap (boys–girls < 3 y) in mean *waz* scores	-0.034	(0.007)	***	-0.024	(0.008)	**	-0.049	(0.008)	***	-0.046	(0.010)	***	0.45	152/67
Gap (girls–boys < 3 y) in % underweight	-1.130	(0.141)	***	-0.867	(0.147)	***	-1.392	(0.167)	***	-1.311	(0.190)	***	0.65	152/67
Gap (men–women 11-15 y) in school attendance	2.547	(0.755)	***	-0.113	(0.801)		2.166	(0.819)	**	0.058	(0.797)		0.29	125/60
Explanatory Variable: AFB, Women 25 – 49 y														
Gap (girls–boys 1–4 y) in risk of mortality	-1.881	(0.408)	***	-1.282	(0.427)	**	-1.416	(0.519)	**	-1.168	(0.519)	*	0.40	185/75
Gap (boys–girls 12–23 m) in % with specified vaccines	-0.519	(0.223)	*	-0.252	(0.245)		-0.089	(0.292)		0.079	(0.292)		0.23	175/73
Gap (boys–girls < 3 y) in mean *haz* scores	-0.023	(0.008)	**	-0.026	(0.009)	**	-0.028	(0.011)	**	-0.038	(0.012)	**	0.24	152/67
Gap (girls–boys < 3 y) in % stunted	-0.525	(0.184)	**	-0.488	(0.200)	*	-0.334	(0.251)		-0.602	(0.274)	*	0.28	152/67
Gap (boys–girls < 3 y) in mean *waz* scores	-0.029	(0.007)	***	-0.027	(0.008)	***	-0.037	(0.010)	***	-0.041	(0.011)	***	0.33	152/67
Gap (girls–boys < 3 y) in % underweight	-0.481	(0.162)	**	-0.385	(0.163)	*	-0.406	(0.219)	!	-0.465	(0.226)	*	0.47	152/67
Gap (men–women 11-15 y) in school attendance	-2.278	(0.757)	**	-0.456	(0.680)		0.281	(0.906)		0.658	(0.733)		0.31	125/60
Controls (of country C, time T, economic change EΔ):		C			C, T			C, EΔ			C, T, EΔ			

NOTE.
Refer to the NOTE to Table 1 for detailed definitions of the variables.
! $p \leq 0.10$, * $p \leq 0.05$, ** $p \leq 0.01$, *** $p \leq 0.001$.

Results excluding India from the sample

Table 4 permits comparison of the results from Column (4) of Tables 2 and 3 with the results for the same models re-run on the sample without India. The bolded and shaded coefficients are ones for which the inferences differed across the two samples. These discrepancies in the inferences support our hypothesis that the effects of fertility decline on gender gaps in children's well-being will depend on initial and prevailing levels of son preference. First, the non-significant associations of fertility decline with girls' nutritional well-being in the full sample became significant in the sample that excluded India. In the latter sample, a decline of one birth in total fertility was associated with increases of 0.06 in girls' mean *haz* and *waz* scores at 0 – 36 months as well as a 1.4 percentage-point reduction in the prevalence of stunting at these ages. In other words, in countries often with lower levels of son preference, girls accrued greater nutritional benefits from fertility decline. Second, whereas fertility decline was associated with a significant *widening* of the gender gap in early childhood mortality in the full sample (reflecting a slower decline in mortality for girls than boys), fertility decline was not associated with changes in the gender gap in mortality in the sample that excluded India. These findings corroborate the idea that an intensification effect of fertility decline may be most pronounced in settings with initially high son preference. Third, whereas increases in the median age at first birth was associated with a significant *decline* in the gender gap in early child mortality in the sample with India (reflecting a faster decline in mortality for girls than boys), an increasing median age at first birth was not associated with changes in the gender gap in mortality in the sample without India. These findings support the idea that later first childbearing accrues equal mortality reductions for boys and girls in less gender-biased societies and that later first childbearing may signal more substantial normative changes in higher son-preference societies, which encourages a reduction in extreme biases against girls.

TABLE 4. Populated-Weighted Random-Effect Estimates of Total Fertility Rate (TFR) and Median Age at First Birth (AFB) on Women's and Girls' Well-being and Gender Gaps in Children's Well-being, Demographic and Health Surveys for 1985-2008

	TFR, women 15–49 y 0–4 y before survey						Median AFB, Women 25 – 49 y					
	All Countries[a]			India Removed[b]			All Countries[a]			India Removed[b]		
	B	(*se*)	*p*	B	(*se*)	*p*	B	(*se*)	*p*	B	(*se*)	*p*
Women's Well-being												
% with any prenatal care, births in prior 3 y	-6.45	(1.63)	***	-6.89	(1.65)	***	5.18	(1.44)	***	4.99	(1.47)	***
% with trained attendant at delivery, births in prior 3 y	-4.23	(1.60)	**	-5.31	(1.58)	***	7.03	(1.31)	***	7.51	(1.30)	***
% with BMI < 18.5, non-pregnant women with birth in prior 3 y	-5.85	(0.97)	***	-2.94	(0.93)	**	-3.83	(1.12)	***	-4.31	(0.79)	***
Girls' Well-being												
Deaths to girls 1–4 y per 1,000 1–4 y	17.59	(1.78)	***	19.45	(1.71)	***	-10.43	(1.95)	***	-11.34	(1.92)	***
% 12–23 m with specified vaccines	-8.31	(1.71)	***	-10.17	(1.59)	***	5.80	(1.46)	***	5.44	(1.43)	***
Mean *haz* score < 3 y	**0.00**	**(0.03)**		**-0.06**	**(0.03)**	*****	0.12	(0.03)	***	0.08	(0.02)	***
% < 3 y stunted	**-0.40**	**(0.74)**		**1.41**	**(0.68)**	*****	-2.09	(0.64)	**	-1.46	(0.57)	**
Mean *waz* score < 3 y	**0.03**	**(0.04)**		**-0.06**	**(0.04)**	**!**	0.20	(0.03)	***	0.15	(0.03)	***
% < 3 y underweight	-1.08	(0.91)		1.13	(0.77)		-3.77	(0.72)	***	-3.04	(0.58)	***
% 11-15 y attending school	-3.83	(1.34)	**	-4.78	(1.40)	***	3.51	(1.26)	**	3.69	(1.29)	**
Gender Gaps in Well-being												
Gap (girls–boys 1–4 y) in risk of mortality	**-1.21**	**(0.60)**	*****	**-0.46**	**(0.53)**		**-1.17**	**(0.52)**	*****	**-0.60**	**(0.45)**	
Gap (boys–girls 12–23 m) in % with specified vaccines	-0.23	(0.32)		-0.12	(0.31)		0.08	(0.29)		0.04	(0.26)	
Gap (boys–girls < 3 y) in mean *haz* scores	-0.05	(0.01)	***	-0.03	(0.01)	*	-0.04	(0.01)	**	-0.03	(0.01)	**
Gap (girls–boys < 3 y) in % stunted	-1.61	(0.24)	***	-1.06	(0.21)	***	-0.60	(0.27)	*	-0.35	(0.20)	!
Gap (boys–girls < 3 y) in mean *waz* scores	-0.05	(0.01)	***	-0.03	(0.01)	**	-0.04	(0.01)	***	-0.04	(0.01)	***
Gap (girls–boys < 3 y) in % underweight	-1.31	(0.19)	***	-0.92	(0.18)	***	-0.46	(0.23)	*	-0.50	(0.17)	**
Gap (men–women 11-15 y) in school attendance	0.06	(0.80)		1.15	(0.80)		0.66	(0.73)		1.08	(0.72)	

NOTE.
Refer to the NOTE to Table 1 for detailed definitions of the variables.
! $p \leq 0.10$, * $p \leq 0.05$, ** $p \leq 0.01$, *** $p \leq 0.001$.
[a] These results are identical to those presented for Model 4 in Tables 2 and 3.
[b] Modeling strategy is identical to that used for Model 4 in Tables 2 and 3, but observations from India have been excluded.

Alternative fits

The results of fitting the fixed-effects models of these same relationships (Appendices 1 and 2) largely mirror those found in Tables 2 and 3. For the models of women's and girls' well-being, Hausman tests indicated a preference for the fixed effect models only for models of vaccination coverage; however, the estimated coefficients for measures of fertility did not differ substantially across the fixed- and random-effects models, and the estimated coefficients for the random-effects models were more conservative (in that they indicated weaker effects of changing fertility). For the models of gender gaps in well-being, the Hausman test revealed a preference for the fixed-effect model only for models of the gender gap in child mortality. For this outcome, the estimated coefficients for both measures of fertility were not significant in the fixed-effects models but were significant in the random-effects models, and the coefficients for total fertility differed in sign across the fixed- and random-effects models.

Overall, these results support our expectations that declines in fertility and increases in the median age at first birth are associated with improvements in women's and girls' well-being. Also, these results suggest that declining fertility and increasing median ages at first birth can have seemingly contradictory effects on gender gaps in mortality and nutrition. We elaborate on these counterintuitive results below.

DISCUSSION

In this paper, we have examined whether national changes in fertility regimes (from high total fertility and an early median age at first birth to lower total fertility and a later median age at first birth) are associate with: (1) aggregate increases in women's and girls' well-being; and (2) changes in gender gaps in children's well-being, conditional on the broader gender-normative environment. Table 5 and the paragraphs below summarize the results of testing our initial hypotheses.

TABLE 5. Summary of Results of Full Random-Effects Models of Well-being Outcomes on the Total Fertility Rate and Median Age at First Birth, Demographic and Health Surveys for 1985-2008

Explanatory variable:	TFR ↓					Median Age at First Birth ↑				
	Girls	**Gaps**		**Hausman**	**Interpretation**	**Girls**	**Gaps**		**Hausman**	**Interpretation**
Outcome		**Fem Adv**	**Male Adv**	**Girls/Gap**			**Fem Adv**	**Male Adv**	**Girls/Gap**	
Women's Well-being										
% with any prenatal care, births in prior 3 y	↑			RE		↑			RE	
% with trained attendant at delivery, births in prior 3 y	↑			RE		↑			RE	
% with BMI < 18.5, non-pregnant women with birth in prior 3 y	↑			RE		↓			RE	
Girls' Well-being										
Deaths to girls 1–4 y per 1,000 1–4 y	↓	↓	↑	RE/FE	boys initially advantaged, (b) intensification	↓	↑	↓	RE/FE	boys initially advantaged, (c) gender parity effect and/or change in norms
% 12–23 m with specified vaccines	↑	NS	NS	FE/FE	(a), equal accrual of benefits	↑	NS	NS	FE/RE	(a), equal accrual of benefits
Mean *haz* score < 3 y	NS	↓	↑	RE/RE	girls initially advantaged, decreasing gap	↑	↑	↓	RE/RE	girls initially advantaged, increasing gap
% < 3 y stunted	NS	↓	↑	RE/RE	girls initially advantaged, decreasing gap	↓	↑	↓	RE/RE	girls initially advantaged, increasing gap
Mean *waz* score < 3 y	NS	↓	↑	RE/RE	girls initially advantaged, decreasing gap	↑	↑	↓	RE/RE	girls initially advantaged, increasing gap
% < 3 y underweight	NS	↓	↑	RE/RE	girls initially advantaged, decreasing gap	↓	↑	↓	RE/RE	girls initially advantaged, increasing gap
% 11-15 y attending school	↑	NS	NS	RE/RE	(a), equal accrual of benefits	↑	NS	NS	RE/RE	(a), equal accrual of benefits

NOTE.

Refer to the NOTE to Table 1 for detailed definitions of the variables.

NS indicates non-significant result

↑↓ indicates direction of effect (↑ indicates widening gender gap, ↓ indicates declining gender gap; note that males are not always assumed to have the initial advantage)

Fem Adv, Male Adv indicates an assumption about whether males or females have the initial advantage in the indicator of well-being

RE, FE indicates whether the results of the Hausman Test supported using a random effects model or fixed effects model respectively

Summary of Findings

Women's well-being: We proposed that national declines in fertility and increases in the median age at first birth would be associated with improvements in women's health because of population-level reductions in maternal nutritional depletion. We tested this hypothesis by examining the relationship between trends in our measures of fertility and trends in the percentage of non-pregnant, recently delivered women who were underweight. Unexpectedly, in fully adjusted random-effects models with the total fertility rate as the explanatory variable, a decline in total fertility was associated with an increasing percentage of women who were underweight. In the fixed-effects models, however, this association was not significant, and the coefficient for total fertility was not robust to alternative specifications (e.g., changed sign across Models 1 – 4, Appendix A). Thus, further investigation of these findings is needed. Otherwise, estimates of the relationship between the median age at first birth and the percentage of non-pregnant, recently delivered women who were underweight were in the expected direction. In fully adjusted models, increases in the median age at first birth were associated with declines in the percentage of these mothers being underweight, and this association was significant in the (preferred) random-effects model.

Both declines in fertility and increases in the median age at first birth were associated with improvements in women's met needs for well-being, as reflected in the percentage with access to prenatal care and a trained attendant at delivery. These associations were significant in the preferred random-effects model, and the directions of these associations were consistent in the fixed-effects models.[20]

Girls' well-being: Our results overwhelmingly supported our hypothesis that aggregate declines in fertility and later first childbearing would be associated with improvements in girls' met needs for well-being, as well as reductions in the burden of harms to girls arising from unmet needs.

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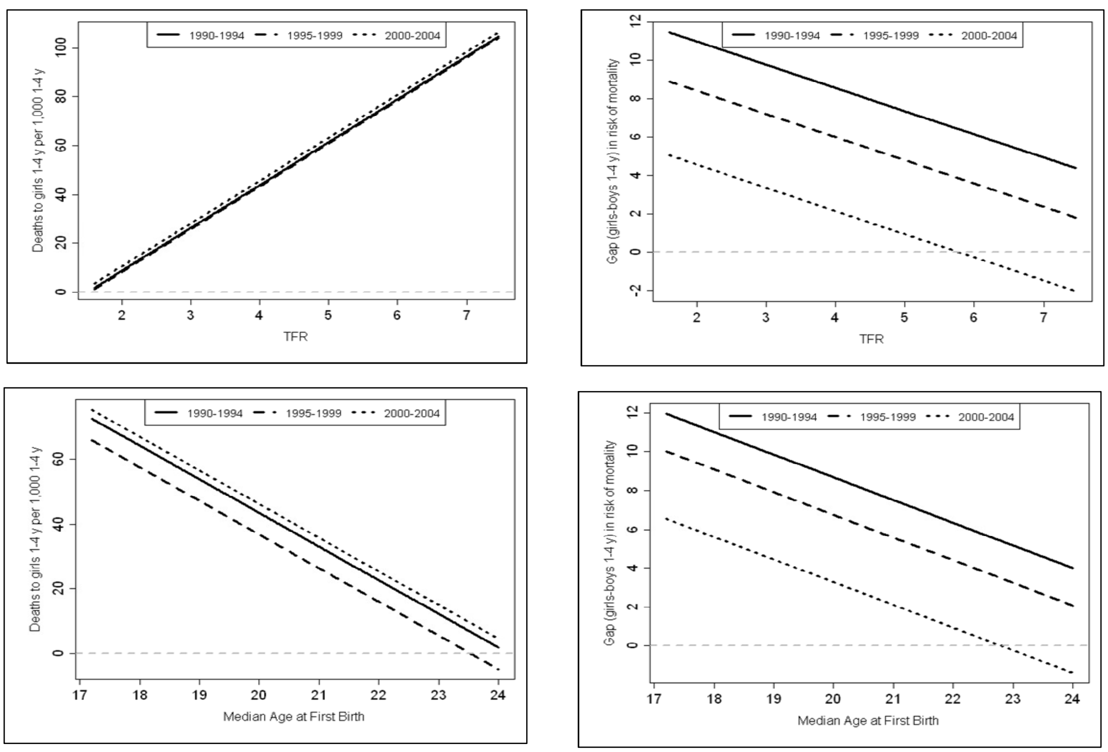
[20] Although the associations of our fertility measures with having a trained attendant at delivery were not significant in the fixed-effects models. As a further sensitivity test, we re-estimated the "full" random-effects models for women's access to maternal health care also controlling for general improvements in health infrastructure by including the percentage of girls fully vaccinated at 12 – 23 months. Adding this control variable weakened the estimated association between our measures of fertility and the measures of women's access to maternal health care; however, all of the relationships remained significant with the exception of that between the total fertility rate and the percentage of births in the prior three years with a skilled attendant at delivery.

Specifically, declines in fertility and increases in the median age at first birth were associated with declines in girls' mortality at 1 – 4 years, increases in their vaccination coverage at 12 – 23 months, and increases in their school attendance at 11 – 15 years. Moreover, a higher median age at first birth was associated with systematic improvements in girls' nutrition at 0 – 36 months. Finally, in models that excluded the case of India, these improvements in girls' well-being were systematically more pronounced in the expected directions. One interpretation of this last finding is that girls benefit more from fertility decline and later first childbearing in countries that are characterized by lower levels of son preference.

Changes in gender gaps in children's well-being: Despite strong evidence that shifting fertility regimes have been associated with improvements in women's and girls' well-being, the relationship between changes in fertility and gender gaps in children's well-being has been more complex.

FIGURE 4. Predicted Trends in Child Mortality of Girls 1-4 y and Gender Gaps in Child Mortality Derived from Results of Random-Effect Models in Tables 2 and 3 (Model 4).



NOTE. Other control variables (socioeconomic conditions) are set to their weighted mean values for the indicated periods.
1990 - 4: GDP per capita $653.6, 33.8% urban, 0.0 per 100 cell phone subscribers, and $11.7 net ODA.
1995 - 9: GDP per capita $825.7, 34.7% urban, 0.6 per 100 cell phone subscribers, and $11.3 net ODA
2000 - 4: GDP per capita $857.3, 38.1% urban, 7.4 per 100 cell phone subscribers, and $17.2 net ODA.

Figure 4 illustrates the seeming discrepancies in our estimated associations of declines in total fertility and increases in the median age at first birth with the gender gap in early childhood mortality. Specifically, these graphs show predicted values for girls' risk of mortality at 1 – 4 years and the gender gap in this risk at specified levels of total fertility and the median age at first birth, using estimated coefficients from the corresponding full random-effects models (Model 4, Tables 2 and 3) and setting the control variables (GDP per capita, percentage urban, mobile cellular subscriptions per 100 people, and net ODA per capita) to their mean values for the periods 1990 – 4, 1995 – 9, and 2000 – 4, respectively. Thus, these predictions illustrate possible relations of changes in fertility with changes in these outcomes for various fixed socioeconomic environments.

As expected, reductions in total fertility and increases in the median age at first birth predicted substantial reductions in girls' mortality. Yet, while reductions in total fertility predicted an *increasing* gender gap in mortality (resulting from initially lower and faster declines in boys' mortality), increases in the median age at first birth predicted a *decrease* in this gender gap. If declines in total fertility and increases in the median age at first birth had occurred simultaneously, our results would be nonsensical because the results for total fertility would suggest an intensification effect (Scenario (b), Figure 2) while the results for the median age at first birth would suggest a gendered parity effect (Scenario (c), Figure 2). However, as we have discussed, age-specific changes in fertility typically have occurred sequentially, with initial declines in total fertility often resulting from restrictions in fertility at older ages and the postponement of first births often occurring later in the fertility transition (e.g., Knodel 1977; Sneeringer 2009). This sequencing of changes in fertility is observable in our sample (Figure 3a).

Thus, if the above interpretations are correct, then the gender gap in mortality may have followed a trajectory similar to that presented in Scenario (d), Figure 2. In the initial stage of fertility decline, women who limited their fertility at older ages had more resources to invest in their children. Both the widening of the gender gap in child mortality (assuming an initial boys' advantage) and the narrowing of the gender gap in measures of nutrition (assuming an initial girls' advantage) indicate that these newly available resources were initially disproportionately invested in boys, resulting in an intensification effect of initial fertility decline. Arguably, these disproportionate investments in boys most likely occurred in the private rather than the public sphere, given that increases in vaccination coverage at 12 – 23 months and school attendance at 11 – 15 years accrued equally to girls and boys over the period (Table 3).

Subsequently, postponement of fertility associated with an increase in the median age at first birth may have signaled a broader shift in gender norms (e.g., McDonald 2000). As the value of women and girls rose, the relative investments in boys and girls shifted, and girls were able to catch up with boys (in the case of the gender gap in mortality) or regain their natural biological advantage (in the case of nutritional status). The results of analyses with the sample that excluded India show that the intensification effects of fertility decline and any catch-up of girls associated with later childbearing are likely to be weaker in lower-son-preference settings.

Strengths and Limitations of the Analysis

Our analysis had some notable strengths. To our knowledge, it is the only panel analysis using the DHS to assess the influences of fertility declines on aspects of women's well-being, girls' well-being, and gender gaps in children's well-being. To conduct this analysis, we have leveraged data from as many as 187 DHS from as many as 75 countries and spanning a period of 23 years. We have linked these data to other information on national socioeconomic conditions for a similar period. Finally, our analyses have made rigorous and systematic efforts to adjust for both fixed- and time-varying sources of confounding in the relationship between fertility declines, women's and girls' well-being, and gender gaps in children's well-being.

Some limitations of our analysis suggest promising avenues for further research. First, while our window of observation spanned 23 years from 1985 to 2008, we were unable to observe the full trajectory of changes in well-being that may arise from changes in fertility. As the DHS continue, the period of observation will be extended, and the potential to conduct time-series analyses of more recently collected data on these and other aspects of women's well-being (such as anemia) will be strengthened.[21] Second, because this was a cross-national time-series study, the results should be interpreted at the aggregate level. A complimentary multilevel analysis might examine how meso-level changes in fertility, desired fertility, and son preference may have affected parental investments in the well-being of sons and daughters. Finally, our analysis lacked direct measures of changes in national norms pertaining to fertility and gender preferences. In the future, researchers might apply our analytical approach to study district-level changes in fertility with our outcomes in

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[21] Using available data in our sample on the percentage of women 15 – 49 years with any anemia (see endnote 9) and full random-effects models analogous to Model 4 in Tables 2 and 3, a decline in fertility and a rising median age at first birth were associated with a decline (albeit not significant at the $p < 0.05$ level) in the prevalence of anemia among women.

countries with sufficiently variable sex ratios at birth and reliable panel data for districts on other relevant variables.

Conclusions and Policy Implications

Together, our findings suggest that fertility decline is associated with improvements in women's and girls' well-being, as reflected by increased access to prenatal care, a trained attendant at birth, child vaccination, and school attendance. Our findings also suggest that fertility decline is associated with broad declines in girls' risks of mortality and malnutrition, especially where son preference has historically been weaker. However, the influences of changing fertility regimes on gender gaps in children's well-being may be varied. On the one hand, fertility decline appears to accrue equal benefits for boys and girls with respect to some intermediate needs for well-being, such as vaccination coverage and school attendance. The concurrent expansion of public health and educational infrastructures may explain some of this relationship, and further research is warranted to explore this possibility. On the other hand, initial fertility declines may have had an intensification effect on gender gaps in child mortality and malnutrition, especially in societies with initially higher son preference. However, later increases in women's median age at first birth, which may signal (especially in more gender-biased settings) shifts toward more egalitarian gender norms, are associated with greater improvements for girls than boys in survival and nutritional status. Thus, the influence of fertility decline on gender gaps in children's well-being (1) may depend on the extent of son preference at the initial stages of fertility decline and (2) may vary across the stages of fertility transition as norms about gender shift concurrently. To avert an initial intensification response to fertility decline, family planning programs in early transition settings might consider efforts to promote gender equity in parental investments in children.

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