Demography (2014) 51:535–561
DOI 10.1007/s13524-014-0282-0

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

Kathryn M. Yount · Sarah Zureick-Brown · Nafisa Halim · Kayla LaVilla

Published online: 25 March 2014

Abstract The influences of recent dramatic declines in fertility on girls' and boys' well-being in poorer countries are understudied. In panels of 67–75 poorer countries, using 152–185 Demographic and Health Surveys spanning 1985–2008, we examined how declining total fertility and women's increasing median age at first birth were associated with changes in girls' well-being and gender gaps in children's well-being, as reflected in their survival, nutrition, and access to preventive healthcare. In adjusted random-effects models, these changes in fertility were associated with gains in girls' survival at ages 1–4 years, vaccination coverage at ages 12–23 months, and nutrition at 0–36 months (for women's later first childbearing). Declining total fertility was associated with similar gains for boys and girls with respect to vaccination coverage but intensified gender gaps in mortality at ages 1–4 years and malnutrition at ages 0–36 months, especially in higher-son-preference populations. Later increases in women's median age at first birth—reflecting more equitable gender norms—were associated with declines in these gaps. Promoting equitable investments in children through family planning programs in higher-fertility societies is warranted.

Keywords Cross-national time series · Demographic and Health Surveys · Fertility trends · Gender gaps in well-being · Girls' well-being

K. M. Yount (✉)
Asa Griggs Candler Chair of Global Health, Hubert Department of Global Health and Department of Sociology, Emory University, 1518 Clifton Rd. NE, Atlanta, GA 30322, USA
e-mail: kyount@emory.edu

S. Zureick-Brown · K. LaVilla
Department of Sociology, Emory University, Atlanta, GA, USA

S. Zureick-Brown
e-mail: Sarah.Zureick-Brown@emory.edu

K. LaVilla
e-mail: kaylalavilla@gmail.com

N. Halim
Department of International Health, Boston University, Boston, MA, USA
e-mail: nhalim@bu.edu

Introduction

The second half of the twentieth century witnessed dramatic declines in fertility, especially in poorer countries (Bongaarts 2008), as well as increasing ages at first birth for women (Frekja et al. 2010; Frejka and Sardon 2006; Westoff 2003; Wilson 2001). For decades, demographers have studied the reasons for these shifts (e.g., Bryant 2007; Hill 1992; Mauldin 1978, 1982; Mauldin and Segal 1988; Tsui and Bogue 1978), including changes in women's opportunities (e.g., Adamchak and Ntseane 1992; Folbre 1983; London 1992; Malhotra et al. 1995; Mason 1984, 1987; Sanderson and Dubrow 2000; Weinberger et al. 1989). Yet, recent dramatic shifts in fertility regimes in poorer countries may have markedly influenced women's lives (Mason 1997; McDonald 2000), including the survival, nutrition, and health care of girls, both overall and relative to boys (Das Gupta and Bhat 1997; Das Gupta and Shuzhuo 1999). These reciprocal influences are understudied in poorer countries.[1]

We examined, over a 23-year period in poorer countries, how declining total fertility and women's increasing median age at first birth have been associated with changes in girls' well-being as well as changing gender gaps in children's well-being. We linked comparable national data from as many as 187 Demographic and Health Surveys (DHS) to other national data sources for as many as 75 poorer countries from 1985 to 2008 (Measure DHS 2011). Underpinning these analyses were three hypotheses:

> *Hypothesis 1 (H1):* Declining total fertility *and* women's later first childbearing will be associated with aggregate improvements in girls' well-being, as reflected by their survival, nutritional status, and vaccination coverage.

> *Hypothesis 2 (H2):* Because equitable gender norms may emerge after lower-fertility norms, boys' well-being should improve faster than girls' well-being as the TFR declines initially, but this pattern should predominate in historically higher-son-preference populations.

> *Hypothesis 3 (H3):* Because increases in women's median age at first birth in the later stages of fertility transition signal more equitable gender norms, girls' well-being should improve faster than boys' well-being as women's median age at first birth increases, especially in historically lower-son-preference societies (Das Gupta and Bhat 1997; Das Gupta and Shuzhuo 1999).

In the upcoming discussion, we clarify our definition of "well-being," the aspects of children's well-being that we analyzed, and the theory underlying our hypotheses. We then describe the sample, data, variables, methods, and results, including (1) levels and trends in our main constructs and (2) focal estimates from random-effects and country-time fixed-effects panel regressions, unadjusted and adjusted for fixed effects and time-varying national socioeconomic conditions. We end by discussing the implications of our findings for theory, research, and policy.

[1] Scholars discuss these reciprocal influences (Brewster and Rindfuss 2000; Huber 1991; Mason 1997; McDonald 2000) and explore them empirically in wealthier Western countries (Engelhardt et al. 2004; Engelhardt and Prskawetz 2004; Kögel 2004; Matysiak and Vignoli 2008).

Human Needs, Their Satisfaction, and Girls' Well-being

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 [enjoy] a satisfactory quality of life" (Economic and Social Research Council 2007:1). Well-being is enabled through investments in *basic needs* for survival, health, and nutrition, which can be satisfied by meeting *intermediate needs*, such as those for adequate preventive healthcare (Gough et al. 2007). The outcomes in this article capture (1) changes in girls' basic needs for survival and nutrition, (2) changing aggregate investments in girls' intermediate needs for vaccination coverage, and (3) changing gender gaps in (1) and (2). Hereafter, we refer to these outcomes as measures of girls' well-being and gender gaps in children's well-being.

Changes in Total Fertility and Women's Median Age at First Birth

To understand how these outcomes may relate to changes in fertility regimes first requires an understanding of global patterns of fertility change. With birth intervals held constant, declines in the total fertility rate (TFR) will occur with restrictions in fertility in the later reproductive years and with increases in women's median age at first birth (AFB), both of which shorten women's potential reproductive life span (Andersson et al. 2009). In Asia, Europe, and Latin America (Andersson et al. 2009; Knodel 1977; Rosero-Bixby et al. 2009), declines in total fertility were characterized initially by restrictions in fertility at the older ages (≥35 years) and subsequently by increases in women's median age at first birth. In Africa, contrary to earlier research (Caldwell et al. 1992), analyses of 74 DHS in 30 sub-Saharan African countries between 1986 and 2006 have shown that "fertility decline [also] began with women later in their reproductive years (age 25 and over), followed by younger age groups" (Sneeringer 2009:21). This series of changes in fertility-related behavior is seen in the sample of countries for this analysis (Fig. 1). Within countries (see bolded examples) and across countries over time (see median cubic spline), the median age at first birth for women aged 25–49 increased most dramatically after the TFR for women aged 25–49 had fallen below 4.0.

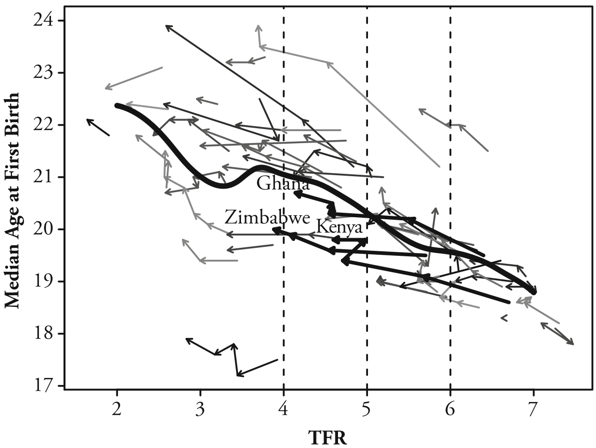

Fig. 1 Trends in total fertility rate, women aged 15–49 at 0–4 years before the survey and median age at first birth, women aged 25–49, Demographic and Health Surveys for 1985–2008

Gender Systems of High- and Low-Fertility Societies

Scholars have asserted that distinct gender systems are inherent in high- and low-fertility regimes and that changes in fertility-related behaviors signal changing gender norms. According to McDonald (2000:431), women's extensive allocation of time and energy to childbearing and child rearing is "implicit in the gender system of a high-fertility society," and declines in fertility imply that a society places less emphasis on this division of labor. Moreover, the impact of fertility decline on women's lives may not occur in the early stages of fertility transition but instead, after a sustained period of decline and a corresponding sustained expectation of improvement in women's lives (McDonald 2000). Other scholars have elaborated that following initial declines in total fertility, increases in women's median age at first birth signal changing norms toward greater equity in women's and men's opportunities (Andersson et al. 2009; Presser 1971; Rosero-Bixby et al. 2009).

Total Fertility, Age at First Birth, and Girls' Well-being

How, then, might initial declines in total fertility and later increases in women's median age at first birth influence the well-being of girls, both overall and relative to boys? To answer this question, we consider gender-*undifferentiated* and gender-*differentiated* models of these relationships.

Gender Undifferentiated Model of Fertility Change and Girls' (Relative) Well-being

A gender-undifferentiated model has the assumption that all children benefit equally from the gains in maternal well-being and resources for investment that accompany initial declines in total fertility and later increases in women's median age at first birth (Fig. 2). With both fertility-related changes, women should experience a lower average lifetime risk of maternal death because they are having fewer births *too early* (before age 20 years) and *too late* (after age 39 years) (Jain 2011; Ross and Blanc 2012). These shifts in fertility also alleviate the reproductive morbidity and maternal *nutritional depletion* that may arise from early, late, and frequent childbirth overlapping with

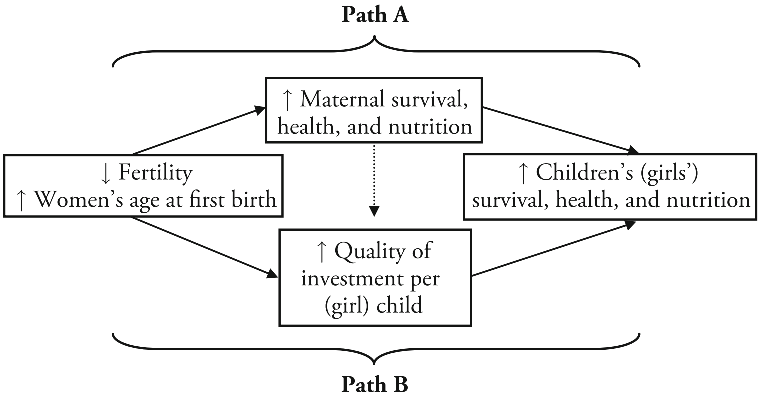

Fig. 2 Hypothesized effects of lower total fertility and women's higher median ages at first birth on girls' well-being: Gender-undifferentiated model

lactation (e.g., King 2003; Merchant and Martorell 1988; Winkvist et al. 1992). In turn, gains in maternal survival, health, and nutrition should improve the survival, health, and nutrition of *all* children (Fig. 2, Path A).

Cross-national and time-series analyses show associations of (changes in) measures of fertility with (changes in) maternal health and survival. Across 79 poorer countries, the TFR has been positively associated with the maternal mortality ratio (MMR), and women's average age at first marriage has been negatively associated with the MMR (Shen and Williamson 1999). From 1990 to 2008, 1.7 million maternal deaths in poorer countries were averted from reductions in the general fertility rate (Ross and Blanc 2012); and in South Asia, 35 % of the 121,000 maternal lives saved were attributed to declining birth rates (Jain 2011). According to a systematic review, birth spacing positively affects maternal anthropometric, micronutrient, and blood hemoglobin status (Conde-Agudelo et al. 2012). In turn, other cross-national and time-series analyses show associations of (changes in) measures of maternal health and survival with (changes in) measures of child health and survival. From 1960 to 1991 in 86 poorer countries, a declining MMR was associated with an increasing probability of child survival to age 5 (Shen and Williamson 1997), and systematic reviews show positive effects of nutrition supplementation in pregnancy on birth outcomes (Haider et al. 2011; Imdad and Bhutta 2011a), child survival (Imdad et al. 2011), and linear growth in childhood (Imdad and Bhutta 2011b).

The second part of this gender-undifferentiated model adapts theories of the quantity-quality tradeoff that occurs with fertility decline (Becker et al. 1960; Becker and Lewis 1973). *Child quality* is a function of the time and resources that parents invest per child (Willis 1973). As the number of births declines, mothers transition from investing widely in many children with poor returns to concentrating investments in fewer children (Reher 2011). According to Lee (2003), women have shifted from spending 70 % of their adulthood on bearing and rearing young children before the demographic transition (in 1800) to spending 14 % today. Emerging norms favoring smaller completed family sizes,[2] and later maternal ages at first birth should free up resources and encourage greater per child investments (Fig. 2, Path B) (Becker et al. 1960; Becker and Lewis 1973; Hanushek 1992). In 19 lower- and higher-income countries between 1994 and 2004, the TFR had negative adjusted associations with per child (public and private) human capital spending (Lee and Mason 2010). In single countries, vaccination is less likely for higher-birth-order children (De and Bhattacharya 2002; Parashar 2005).

Cross-national time-series analyses also have captured the "total" adjusted associations of changes in fertility with changes in child mortality and nutrition (Paths A + B, Fig. 2). Based on data for 66 urban and rural areas in 33 lower-income countries with two DHS between 1986 and 1998, the probabilities of dying at various ages below 1 year rose for infants of mothers older than 35 with birth intervals of less than 24 months (Rutstein 2000). From 1965 to 1991, in 55–70 poorer countries, a higher contraceptive prevalence was associated with a lower rate of infant mortality (Shen and Williamson 2001). In a multilevel analysis of DHS data from 42 countries since 1990, children age 5 or younger had lower probabilities of dying in infancy or of being stunted[3] among mothers with longer birth intervals (Heaton et al. 2005). Lastly, in a multilevel analysis

[2] The total number of births a woman has in her reproductive lifetime.

[3] Children's height-for-age *z* scores are more than 2 standard deviations below the median height-for-age *z* scores in the NCHS/CDC/WHO reference population (Rutstein and Rojas 2006).

of DHS data from 35 countries since 2000, community-level measures of *reproductive norms* (mean age at first union, mean birth interval) were positively associated with height-for-age *z* scores in children younger than age 5 (Fox and Heaton 2012). Thus, from this perspective, as the TFR declines and women's median AFB rises, aggregate gains in maternal well-being and resources for investment should similarly enhance the well-being of girls and boys.

Gender-Differentiated Model of Fertility Change and Girls' (Relative) Well-being

According to a gender-differentiated model of fertility change and children's well-being, the *relative* benefits to girls versus boys from *pro-quality norms* that emerge with declining fertility will depend on prevailing gender norms (e.g., Das Gupta and Bhat 1997; Das Gupta and Shuzhuo 1999). If resource disadvantages accrue to girls at higher parities (e.g., Das Gupta 1987; Muhuri and Preston 1991; Pebley and Amin 1991), then declining total fertility should improve girls' relative well-being by eliminating the parities at which gender biases emerge. Das Gupta and Bhat (1997) called this pattern a "parity effect," but we call it a "gendered parity effect" to distinguish it from the equitable gains in children's well-being that may arise from declining total fertility (described earlier). Alternatively, declining total fertility may spur an *intensification* of gender gaps in children's well-being favoring boys because existing son preference exerts "pressure" at each parity to invest in boys.

Das Gupta and Bhat (1997) theorized how declining *total* fertility alters gender gaps in children's well-being under conditions of high son preference or a lag between lower-fertility norms and more equitable gender norms. We consider also the potential effects of women's increasing median age at first birth, which itself is a sign of more equitable gender norms. Scenarios (a)–(d) in Fig. 3 depict how these changes in fertility may affect trajectories in gender gaps in children's well-being. In all cases, an initial gender gap in well-being favoring boys is assumed.

Stable gender gap. Scenario (a) shows balanced gains in well-being for boys and girls, resulting in a stable gender gap in well-being. In populations that initially prefer gender-equitable investments, boys and girls should benefit similarly from declining fertility and increasing median ages at first birth (barring compensatory investments for past inequities under resource constraints). In populations where son preference manifests at the higher parities and lower-fertility norms precede more-equitable gender norms, Scenario (a) reflects a balanced and persistent influence of the gendered-parity and intensification effects.

Predominant intensification effect. Scenario (b) reflects persistent conditions of non-parity-specific high son preference or a lag between lower-fertility norms and equitable gender norms (Das Gupta and Bhat 1997; Basu 1999). As total fertility falls, these normative conditions create pressure at each parity to invest in sons, and thus boys reap the benefits of declining fertility more quickly than do girls. In India during the 1980s, total fertility fell by 20 %, but the number of sons desired by women who had none fell by only 7.4 % (Registrar General of India 1981, 1991; Operations Research Group 1980, 1988; both cited in Das Gupta and Bhat 1997). An intensification effect has been observed in China and India, where girls' excess mortality at a given parity has grown

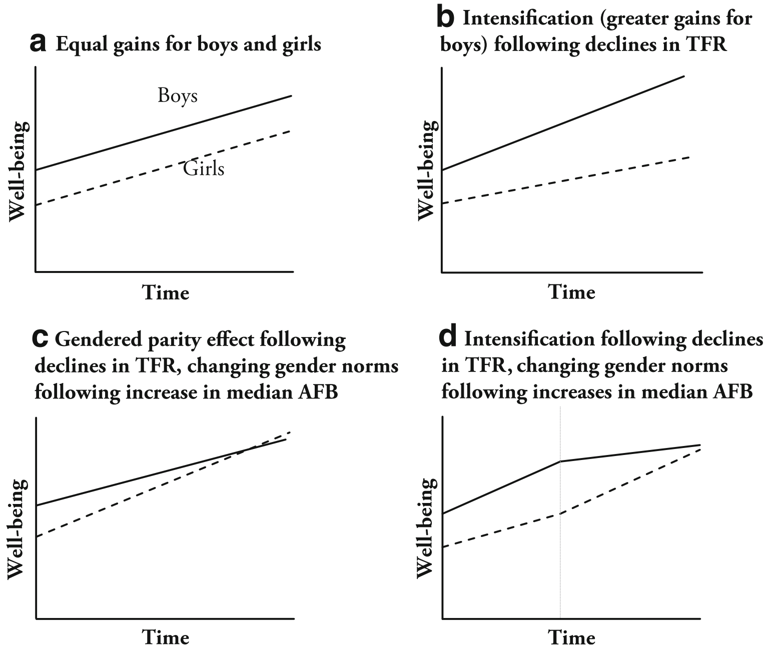

Fig. 3 Four scenarios depicting likely trajectories of gender gaps in well-being following changes in fertility. Adapted from Das Gupta and Bhat (1997)

as total fertility has fallen (Das Gupta and Bhat 1997; Hull 1990; Zeng et al. 1993). Because son preference is more common than daughter preference in poorer countries (Fuse 2010; World Values Survey Association 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. Negative associations between total fertility and the male-to-female sex ratio at birth across 148 countries (Barber 2004), China (Croll 2000; Junhong 2001), five states in India (Basu 1999; Dyson 2001; Vella 2005),[4] and South Korea (Croll 2000) support this interpretation.

Predominant gendered parity effect. Scenario (c) depicts the persistent situation in which girls benefit more quickly from declining fertility than do boys. In higher-son-preference populations, this effect may arise because the gendered parity effect outweighs the intensification effect. More likely, this scenario arises because declining total fertility (especially from later increases in women's median ages at first birth) coincides with shifting gender norms toward a higher valuation of women or girls (Andersson et al. 2009; McDonald 2000; Presser 1971; Rosero-Bixby et al. 2009). In 93 countries, lower fertility was associated with a rising male-to-female infant mortality rate ratio toward levels of 115 to 130, which are the expected ratios given male infants' greater biological vulnerability (Fuse and Crenshaw 2006). The crossover of the gender gap in well-being in Scenario (c) shows how changing gender norms may foster girls'

[4] These states are Gujarat, Haryana, Maharashtra, Punjab, and Tamil Nadu.

long-run advantage, at least for some aspects of well-being such as grade completion (Knodel 1997; Grant and Behrman 2010).

Intensification followed by changing gender norms. Scenario (d) depicts the combination of Scenario (b) followed by Scenario (c). Here, boys benefit more than girls initially because early declines in fertility occur under son preference or a lag between lower-fertility norms and more equitable gender norms (*intensification*, Scenario (b)). As women's median age at first birth rises and signals more equitable gender norms, girls experience faster gains in well-being than do boys (Scenario (c)). In cross-regional analyses of national surveys and vital-registration data, the male-to-female ratio of mortality at ages 1–4 in Southern Asia declined and then rebounded as overall child mortality fell (Sawyer 2012).[5] A similar trend in the sex ratio at birth occurred in South Korea but was attributed to industrialization and urbanization (rather than to initial fertility decline followed by women's rising median age at first birth) (Chung and Das Gupta 2007).

Three expectations follow from this discussion. First, declining total fertility *and* women's later first childbearing will be associated with aggregate improvements in girls' well-being, as reflected in their improved survival, nutritional status, and vaccination coverage (H1). Second, *relative* gains in the well-being of girls versus boys will depend on prevailing gender norms and the stage of fertility decline. Namely, because more equitable gender norms may emerge after lower-fertility norms, the well-being of boys will improve faster than that of girls as the TFR initially declines, but this pattern will predominate in historically higher-son-preference populations (H2). Third, because an increase in women's median age at first birth in the later stages of fertility decline itself is one sign of more equitable gender norms, the well-being of girls will improve faster than that of boys as women's median age at first birth increases, especially in historically lower-son-preference populations (H3) (Das Gupta and Bhat 1997; Das Gupta and Shuzhuo 1999).

Sample and Data

The country was the unit of analysis, and the sample included unequal panels of 60–75 countries in which 152–185 DHS was conducted from 1985 to 2008.[6] With core funding from the U.S. Agency for International Development, the DHS are national household surveys conducted about every five years in lower- and middle-income countries in sub-Saharan Africa, North Africa/West Asia/Europe, Central Asia, South and Southeast Asia, and Latin America and the Caribbean. In all countries, household heads receive standardized questions on member demographics and household economic status, and women aged 15–49 receive standardized questions on maternal and child health and nutrition, reproductive health, and fertility preferences.

[5] This ratio remained below 80 per 100 into the 2000s.

[6] For countries in earlier but not later DHS rounds, initial estimations of fixed-effects models with data on selected outcomes imputed from the Multiple Indicator Cluster Surveys (MICS) (UNICEF 2011) showed no meaningful differences in inferences.

Our outcomes, derived from the DHS (Measure DHS 2011), included two sets of variables capturing (1) *the well-being of girls* and (2) *the well-being of girls relative to that of boys* in each survey. For *girls' well-being*, one measure of mortality captured the number of deaths to girls at ages 1–4 years per 1,000 girls aged 1–4 years in the 10 years before the survey.[7] Two measures of nutrition captured the mean height-for-age (*haz*) and weight-for-age (*waz*) *z* scores for girls aged 0–36 months with respect to the WHO/CDC/NCHS international reference population (Rutstein and Rojas 2006). Two other measures of nutrition for girls aged 0–36 months captured the percentage *stunted*, or below –2 standard deviations (SD) from the mean *haz* in the reference; and the percentage *underweight*, or below –2 SD from the mean *waz* score in the reference. One measure for access to health services captured the percentage of girls aged 12–23 months receiving all recommended vaccines (BCG, DPT 1–3, polio 0–3, and measles, according to the child's vaccination card or mother's report).

Measures of *gender gaps in well-being* captured the difference between girls and boys in their risks of mortality at ages 1–4 years and potential nutritional and preventive-care correlates of this gap (e.g., Hill and Upchurch 1995; Pande 2003): the *haz* and *waz* scores at 0–36 months, percentages stunted and wasted at 0–36 months, and percentages fully vaccinated at 12–23 months. 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 at ages 1–4 and for *boys'* minus *girls' haz* scores at 0–36 months reflected disadvantages for girls).

Explanatory variables were derived from the DHS and included two aggregate measures of the fertility regime. The TFR, or total number of births per woman of reproductive age, was estimated from age-specific fertility rates for women aged 15–49 in the period 0–4 years before each DHS. The median age at first birth was taken for women aged 25–49 at each DHS.

Two *covariates* were 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) across the years of each DHS for a given country. Four other time-varying covariates captured socioeconomic changes that may have been correlated with changes in aggregate fertility and child well-being (e.g., Chung and Das Gupta 2007; Shen and Williamson 2001): mobile cellular subscriptions per 100 people in the same year as each DHS; percentage of the population living in urban areas in the same year as or up to six years before each DHS; and the average gross domestic product (GDP) per capita and net Official Development Assistance (ODA) per capita received in the period 0–4 years before the DHS (in current U.S. dollars). The data for these variables came from the World Development Indicators database (World Bank 2011).

Methods

Descriptive analyses included univariate distributions of all variables to assess their completeness and distributions, within-country trends in all variables to

[7] A measure for early childhood mortality averaged over a shorter interval of time was not available from the online national statistics database for the DHS.

ensure sufficient change over time to permit time-series analyses, and bivariate plots of all outcomes and explanatory variables to explore potential nonlinearities in their associations.

Multivariate models for each outcome were estimated using fixed- and random-effects approaches. In general, fixed- and random-effects models can be used to explore the relation between an explanatory variable and an outcome in panel data. Fixed-effects approaches involve modeling these relationships within entities such as countries, thereby controlling for country-specific time-invariant attributes that may confound the estimated relationships of interest. These models also assume that time-invariant attributes are unique to a given country, so the country's error term and the constant capturing the country's attributes should not be correlated with those of other countries. Random-effects models do not impose this assumption but instead assume that differences across countries are random and uncorrelated with either the explanatory or outcome variables. The equations for the country random-effects models that were estimated with the TFR and the median AFB as explanatory variables are, respectively

$$Y_{it} = \beta_{TFR} TFR_{i(t,t-4)} + \alpha_{TFR} + u_i + e_{it} \tag{1a}$$

$$Y_{it} = \beta_{AFB} AFB_{it} + \alpha_{AFB} + u_i + e_{it}, \tag{1b}$$

where Y_{it} is the dependent variable in country i ($i = 1, \ldots, N$) at time t ($t = 1, \ldots, T$). $TFR_{i(t,t-4)}$ and AFB_{it} are the explanatory variables for country i during the period 0–4 years before the survey (for the TFR) or at the time of the survey (for the AFB). β_{TFR} and β_{AFB} are the respective coefficients for these explanatory variables, and α_{TFR} and α_{AFB} are the respective unknown intercepts. The u_i are between-country errors, and the e_{it} are within-country errors. Two other models then were estimated that adjusted for time fixed effects and national socioeconomic changes:

$$Y_{it} = \beta_{TFR} TFR_{i(t,t-4)} + \beta_{\mathbf{Z}} \mathbf{Z}_t + \beta_{\mathbf{X}} \mathbf{X}_{it-} + \alpha_{TFR} + u_i + e_{it} \tag{2a}$$

$$Y_{it} = \beta_{AFB} AFB_{it} + \beta_{\mathbf{Z}} \mathbf{Z}_t + \beta_{\mathbf{X}} \mathbf{X}_{it-} + \alpha_{AFB} + u_i + e_{it}, \tag{2a}$$

where $\mathbf{Z}$ is a vector of $T - 1$ calendar years entered as design variables, and $\mathbf{X}_{it-}$ is a vector of national socioeconomic conditions for country i at time $t-$ (with each interval described in the earlier variable section). Interpretations of the coefficients β_{TFR} and β_{AFB} in random-effects models include the within-country and between-country effects and represent the average effect of the TFR (or AFB) over Y when the TFR (or AFB) changes across time and between countries by one unit.

In the aforementioned models, the population sizes for each country, averaged over the period of analysis, were used as analytic weights (Dorius 2008).[8] Decisions about data weights are important considerations in cross-national analyses (e.g., Wilson 2001; Dorius 2008; Dorius and Firebaugh 2010). Some researchers weight countries equally

[8] Population estimates from the 2008 revision of *World Population Prospects* were matched by year to each DHS (United Nations 2009) and averaged within country to create country-specific weights.

if the units of interest are *economies* or *states* (e.g., Klasen and Lamanna 2009). This approach would be useful to ask whether a gender gap in well-being is rising or falling in the *average country*, without accounting for each country's population size (Dorius and Firebaugh 2010). Yet, if the interest lies in the welfare of *populations* (composed of individuals), sociologists have argued that the suitable approach is to weight countries by population size (e.g., Firebaugh 1999, 2003; Goesling and Firebaugh 2004; Korzeniewicz and Moran 1997). This decision ensures that a change in *Y* for a more populous country, such as India, has a greater influence on the trend in inequality than a change in *Y* for less populous country, such as Bhutan. Because our focus is on changes in both *girls'* well-being and gender gaps in *children's* well-being, all national estimates in our analyses are weighted by their average population size for the period of study.[9]

Several diagnostic tests were performed to assess the fits of models reflected in Eqs. (2a) and (2b). First, *F* tests for the joint significance of the design variables for calendar year confirmed that their inclusion was warranted. Second, variance inflation factors (VIFs) estimated from ordinary least squares (OLS) regressions controlling for the socioeconomic indicators $\mathbf{X}_{it}$ and time fixed effects (1.2 to 6.0) suggested that multicollinearity was not influencing the least square estimates (Hair et al. 1995; Marquardt 1970; Neter et al. 1989; O'Brien 2007). Third, Hausman tests were performed to assess the relative fits of fixed- versus random-effects models that did not include average population sizes as analytic weights (Hausman 1978).[10] Significant test statistics would imply that the fixed-effects models were preferred; yet, only 4 of the 24 Hausman tests were significant at $p \leq .05$, and in only one of these 4 cases did the inference for total fertility differ. Thus, the random-effects estimates are presented as main results, and the fixed-effects estimates are discussed and available upon request. Fourth, alternative random-effects models were specified to assess the sensitivity of the findings to (1) including quadratic terms for measures of fertility, (2) using age-specific fertility rates to capture fertility declines at younger (20–24 years) and older (40–44 years) ages, (3) using the median age at first birth for women aged 30–34 at each DHS, and (4) adding interaction terms between the main measures of fertility and measures for region, baseline son preference, and socioeconomic conditions. These results are summarized and available upon request.

Finally, extending prior cross-national studies of gender gaps in child mortality (e.g., Sawyer 2012), we reran all final random-effects models excluding the observations for (1) India and (2) the five most populous *high-son-preference* countries in the sample, in which at least 20 % of women in the DHS reported a preference for sons (India 22.1 %, Pakistan 35.3 %, Nigeria 24.4 %, Ethiopia 22.1 %, and Democratic Republic of Congo 26.9 %, in order of population size according to average population weights in our analysis) (Fuse 2010). Excluding India alone had two motivations. First, India was most

[9] For measures of girls' well-being, the coefficients in weighted and unweighted models were similar in magnitude and significance. For measures of gender gaps in well-being, the associations of TFR were broadly similar across weighted and unweighted models, but those for AFB were attenuated toward zero and lost significance in the unweighted models.

[10] Robust standard errors were estimated for all fixed-effect models to account for country-level clustering (Dorius 2008). Robust standard errors were not estimable for random-effects models with population-average weights. Inferences from unweighted random-effects models with and without robust standard errors and from weighted random-effects models were comparable.

heavily weighted in the full-sample analysis because its average population size for the period of observation (987 million) was five times larger than that for the country that was the next most populous (Indonesia). Second, this exclusion enabled us to assess changes in the coefficients for TFR when a high-son-preference country was excluded (addressing H2 and H3). Although India shows considerable regional variation in levels of son preference (e.g., Dyson and Moore 1983; Dyson 2012), it ranks eleventh among 50 other countries with DHS data in the percentage of women who report son preference (22.1 %) (Fuse 2010), and it has the largest population among the 13 high-son-preference countries in our sample. Our second strategy to exclude observations from the five most populous high-son-preference countries permitted us to explore the robustness of the findings based on excluding India alone. In other words, does any intensification of gender gaps in children's well-being that may occur with declines in total fertility diminish when the largest high-son-preference populations are removed (H2)? Likewise, is any catch-up in girls' versus boys' well-being that may occur with increases in women's median age at first birth accentuated when high-son-preference populations are removed (H3)?

Results

National Characteristics for 1990–1994 and 2005–2008

Table 1 shows population-weighted statistics for all variables for countries with DHS undertaken in 1990–1994 and 2005–2008 as well as for the total sample.

Girls' risk of dying in early childhood averaged 44 deaths per 1,000 in 1990–1994 and 29 per 1,000 in 2005–2008. The mean risk of dying in early childhood was greater for girls than boys in both periods, by 6 deaths per 1,000 in 1990–1994 but by 4 per 1,000 in 2005–2008. The average percentage of girls aged 12–23 months receiving all recommended vaccinations rose from 44 % in 1990–1994 to 51 % in 2005–2008, but the gender gap in vaccination coverage consistently favored boys by about 2 % in both periods. High mean percentages of girls aged 0–36 months were stunted (40 %) and underweight (40 %) in 1990–1994; however, these mean percentages were lower by 1 % to 2 % for girls than for boys in this period. In 2005–2008, the mean percentages of girls who were stunted and underweight were lower (34 % and 37 %, respectively) than in 1990–1994, 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 2005–2008).

The average TFR was 3.9 in 1990–1994 and was one-half birth lower (3.4) in 2005–2008. The average median age at first birth for women aged 25–49 was 20.0 years in 1990–1994 and 20.3 years in 2005–2008. The average number of mobile cellular subscriptions per 100 people grew rapidly from 0 in 1990–1994 to 21 in 2005–2008, as did the average GDP per capita (from $654 in 1990–1994 to $782 in 2005–2008). The average percentage of the population living in urban areas and the average net ODA per capita remained fairly constant between 1990–1994 and 2005–2008, at 34 % and $12, respectively.

Table 1 Descriptive statistics for measures of girls' well-being, gender gaps in children's well-being, fertility, and socioeconomic conditions, Demographic and Health Surveys (DHS) for 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]
Girls' Well-being																	
Deaths to girls 1–4 yrs. per 1,000 girls 1–4 yrs.	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 months with specified vaccines[d]	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 yrs.[e]	–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 yrs. stunted[f]	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 yrs.[g]	–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 yrs. underweight[h]	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
Gender Gaps in Children's Well-being (a positive value reflects girls' disadvantage)																	
Gap (girls–boys 1–4 yrs.) 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 mo.) 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 yrs.) in mean *haz* scores	–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 yrs.) in % stunted	–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 yrs.) in mean *waz* scores	–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 yrs.) in % underweight	–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
National Fertility Regime																	
TFR, women aged 15–49 at 0–4 yrs. 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

Table 1 (continued)

	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
Median age at first birth, women aged 25–49	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 Socioeconomic Indicators																	
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 yrs. prior, current US$[i]	653.6	661.7	188.9	2,613.5	31	781.5	493.1	131.4	3,227.7	39	789.9	760.7	131.4	5,566.0	187	75	2.5
Urban population as a % of total, current, or prior year to DHS	33.8	15.3	5.4	74.8	31	34.4	12.3	12.5	78.3	39	35.7	16.0	5.2	80.1	187	75	2.0
Avg. net ODA received per capita 0–4 yrs. prior, current US$[j]	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

[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] Specific vaccinations include: BCG, DPT 1–3, polio 0–3, and measles.

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

[f] % stunted refers to the percentage below –2 SD from the median *haz* in the WHO/CDC/NCHS international reference population.

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

[h] % underweight refers to the percentage below –2 SD from the median *waz* score in the WHO/CDC/NCHS international reference population.

[i] GDP refers to gross domestic product.

[j] ODA refers to overseas development assistance.

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

Table 2 presents the results of random-effects models exploring the associations of changes in total fertility and women's median age at first birth with changes in girls' well-being and gender gaps in children's well-being. Columns 1a and 1b show the results for models in which only the country is controlled (Eqs. (1a) and (1b)). Columns 2a and 2b show the results for models in which controls for time and socioeconomic changes are added (Eqs. (2a) and (2b)). Comparing coefficients across columns 1a, 1b, 2a, and 2b reveals that associations of the TFR and AFB with girls' well-being were attenuated when time and changing socioeconomic conditions were controlled, whereas most associations of the TFR and AFB with gender gaps in girls' well-being were strengthened when time and socioeconomic changes were controlled. Table 3 synthesizes the findings shown in Table 2.

Girls' Well-being

A declining TFR and women's increasing median AFB both were associated with substantial gains in some measures of girls' well-being (Table 2, columns 2a and 2b). A decline of one birth in the TFR and an increase of one year in women's median AFB were associated, respectively, with 17.6 and 10.4 fewer deaths per 1,000 girls aged 1–4 and increases of 8.3 and 5.8 percentage points in rates of full vaccination coverage at ages 12–23 months. Notably, a decline in the TFR was not associated with changes in girls' nutrition, but a one-year increase in women's median AFB was associated, respectively, with increases of 0.12 and 0.20 in girls' mean *haz* and *waz* scores at ages 0–36 months as well as declines of 2.1 and 3.8 percentage points in girls' rates of stunting and underweight at these ages.

Gender Gaps in Children's Well-being

In turn, a decline in the TFR was associated with an *increase* in boys' advantage in early-childhood mortality (initially favoring boys in 41 of 75 countries), no significant change in the gender gap in vaccination coverage (initially favoring boys in 45 of 73 countries), and a *reduction* or *reversal* of girls' advantage in measures of nutrition at ages 0–36 months (initially favoring girls in 46–64 of 67 countries) (Table 2, column 2a; Table 3). Thus, declines in the TFR were associated with faster declines in early-childhood mortality for boys than girls (intensification), similar increases in vaccination coverage for boys and girls, and improvements in measures of nutrition for boys but not girls. The latter associations with gender gaps in nutrition may be interpreted as intensification, given that the average gender gaps in stunting (–1.4) and wasting (–0.5) for children aged 0–36 months in the reference population reflect a consistent advantage for girls (Cogill 2003).

In contrast, an increase in women's median AFB was associated with a *reduction* or *reversal* of boys' advantage in early-childhood mortality, no change in the gender gap in vaccination coverage, and an *increase* in girls' advantage in measures of nutrition at ages 0–36 months (Table 2, column 2b). Thus, increases in the median AFB were

Table 2 Populated-weighted random-effect estimates of total fertility rate (TFR) and women's median age at first birth (AFB) on girls' well-being and gender gaps in children's well-being, Demographic and Health Surveys for 1985–2008

	TFR, Women Aged 15–49 at 0–4 Years Before Survey								Median AFB, Women Aged 25–49							
	(1a)			(2a)					(1b)			(2b)				
	B	(SE)	*p*	B	(SE)	*p*	R^2	*N*/*n*	B	(SE)	*p*	B	(SE)	*p*	R^2	*N*/*n*
Girls' Well-being																
Deaths to girls 1–4 yrs. per 1,000 1–4 yrs.	18.94	(1.29)	***	17.59	(1.78)	***	0.69	184/74	–15.65	(1.58)	***	–10.43	(1.95)	***	0.31	185/75
% 12–23 months with specified vaccines	–10.35	(1.28)	***	–8.31	(1.71)	***	0.27	176/73	7.39	(1.26)	***	5.80	(1.46)	***	0.24	175/73
Mean *haz* score <3 yrs.	–0.14	(0.03)	***	0.00	(0.03)		0.66	152/67	0.26	(0.03)	***	0.12	(0.03)	***	0.70	152/67
% <3 yrs. stunted	3.14	(0.79)	***	–0.40	(0.74)		0.73	152/67	–5.99	(0.71)	***	–2.09	(0.64)	**	0.75	152/67
Mean *waz* score <3 yrs.	–0.17	(0.03)	***	0.03	(0.04)		0.54	152/67	0.27	(0.03)	***	0.20	(0.03)	***	0.62	152/67
% < 3 yrs. underweight	3.82	(0.61)	***	–1.08	(0.91)		0.61	152/67	–5.78	(0.65)	***	–3.77	(0.72)	***	0.67	152/67
Gender Gaps in Children's Well-being																
Gap (girls–boys 1–4 yrs.) in risk of mortality	–0.757	(0.487)		–1.211	(0.603)	*	0.45	184/74	–1.881	(0.408)	***	–1.168	(0.519)	*	0.40	185/75
Gap (boys–girls 12–23 months) in % with specified vaccines	0.047	(0.258)		–0.229	(0.324)		0.24	176/73	–0.519	(0.223)	*	0.079	(0.292)		0.23	175/73
Gap (boys–girls <3 yrs.) in mean *haz* scores	–0.035	(0.007)	***	–0.048	(0.011)	***	0.47	152/67	–0.023	(0.008)	**	–0.038	(0.012)	**	0.24	152/67
Gap (girls–boys <3 yrs.) in % stunted	–1.135	(0.165)	***	–1.615	(0.236)	***	0.62	152/67	–0.525	(0.184)	**	–0.602	(0.274)	*	0.28	152/67
Gap (boys–girls <3 yrs.) in mean *waz* scores	–0.034	(0.007)	***	–0.046	(0.010)	***	0.45	152/67	–0.029	(0.007)	***	–0.041	(0.011)	***	0.33	152/67

Table 2 (continued)

	TFR, Women Aged 15–49 at 0–4 Years Before Survey								Median AFB, Women Aged 25–49							
	(1a)			(2a)					(1b)			(2b)				
	B	(SE)	p	B	(SE)	p	R^2	N/n	B	(SE)	p	B	(SE)	p	R^2	N/n
Gap (girls–boys <3 yrs.) in % underweight	–1.130	(0.141)	***	–1.311	(0.190)	***	0.65	152/67	–0.481	(0.162)	**	–0.465	(0.226)	*	0.47	152/67
Controls (country C, time Z, socioeconomic indicators (S)[a]		C			C, Z, S					C			C, Z, S			

Note: Refer to the footnotes to Table 1 for detailed definitions of the variables.

[a] Socioeconomic indicators include mobile cellular subscriptions per 100 population, same year as DHS; Avg. GDP per capita 0–4 years prior, current US$; Urban population as a % of total, current, or prior year to DHS; and Avg. net ODA received per capita 0–4 years prior, current US$.

*$p \leq .05$; **$p \leq .01$; ***$p \leq .001$

Table 3 Summary of results of full random-effects models of girls' well-being and gender gaps in children's well-being on the total fertility rate and women's median age at first birth, Demographic and Health Surveys for 1985–2008

Explanatory Variable:	TFR ↓				Women's Median Age at First Birth ↑		
Observed Changes From Baseline:							
Outcome	Girls Initially Disadvantaged (No./Total)[a]	Δ in Girls' Outcome	Δ in Gap	Interpretation	Δ in Girls' Outcome	Δ in Gap	Interpretation
Deaths to Girls 1–4 Yrs. per 1,000 1–4 Yrs.	41/75	↓	↑	(b) Intensification, boys' initial advantage increased	↓	↓	(c) Gender parity effect and/or change in norms; boys' initial advantage reduced or reversed
% 12–23 Months With Specified Vaccines	45/73	↑	NS	(a) Girls and boys benefit similarly; boys' slight overall advantage sustained	↑	NS	(a) Girls and boys benefit similarly; boys' slight overall advantage sustained
Mean *haz* Score <3 Yrs.	3/67	NS	↓	(b) Intensification, girls' initial advantage reduced or reversed	↑	↑	Girls' initial advantage increased
% <3 Yrs. Stunted	9/67	NS	↓	"	↓	↑	"
Mean *waz* Score <3 Yrs.	3/67	NS	↓	"	↑	↑	"
% <3 Yrs. Underweight	21/67	NS	↓	"	↓	↑	"

Notes: Refer to the footnotes to Table 1 for detailed definitions of the variables. NS indicates a nonsignificant association of the fertility-related variable with the outcome. ↑↓ indicates direction of association (↑ indicates widening gender gap, and ↓indicates declining gender gap; empirically, boys do not always have the initial advantage).

[a] Proportion of countries where girls are initially disadvantaged on the indicator of well-being.

associated with faster declines in early-childhood mortality for girls than boys (a gendered parity effect or more likely evidence of more equitable gender norms), similar gains in vaccination coverage for boys and girls, and faster improvements in girls' than in boys' nutrition (Table 3).

Alternative Estimation Strategy and Alternative Model Specifications

The results from fixed-effects models of these same relationships (available upon request) largely mirror those discussed earlier. For girls' well-being, Hausman tests indicated a preference for the fixed-effect models for vaccination coverage only; 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 (indicating weaker influences of changing fertility). For gender gaps in well-being, the Hausman test revealed a preference for the fixed-effect model only for the gap in early-childhood 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, the fixed-effect results supported our expectations that declining fertility and women's increasing median age at first birth were associated with gains in girls' well-being and were associated in contradictory ways with gender gaps in mortality and nutrition.

Alternative model specifications further confirmed our main findings (available upon request). First, additional quadratic terms for measures of fertility were significant in only 1 of 12 adjusted models with women's median AFB and in 7 of 12 adjusted models with the TFR. In the latter seven models, the quadratic terms suggested that the gains in well-being associated with declining fertility were greater at lower levels of fertility. Second, replacing our main measures of fertility with fertility rates at older (e.g., 40–44 years) and younger (e.g., 20–24 years) ages suggested that the intensification of gender gaps in children's early-childhood mortality were associated only with fertility restriction at older ages and that improvements in girls' nutrition were associated with fertility restriction at younger rather than older ages. Third, using the median AFB for women aged 30–34 yielded similar inferences to models using the median AFB for women aged 25–49, although the relationships with measures of girls' well-being were generally attenuated. Fourth, interactions of total fertility and women's median age at first birth with a regional indicator for sub-Saharan Africa (1 = yes, 0 = no) (Kögel 2004)[11] were significant in adjusted models for only a subset of the 12 outcomes (9 and 4, respectively). In these cases, the benefits of fertility decline and later childbearing tended to be weaker in sub-Saharan Africa. Fifth, interactions with an indicator for any son preference at baseline (sex ratio at birth > 1.05 vs. ≤ 1.05 at the time of the first DHS)[12] were not significant in 9 of the 12 adjusted models with the TFR or in any of the 12 adjusted models with women's median AFB. Interactions with the socioeconomic variables showed that the benefits of declining fertility for some aspects of girls' 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 that

[11] This demarcation also arguably captures different regional patterns and timings of fertility change.
[12] Of the observations in this analysis, 18 % had initial sex ratios greater than 1.05.

the benefits of women's later first childbearing for girls relative to boys tended to be weaker with increasing net ODA per capita.

Results Excluding the Largest High-Son-Preference National Populations

Table 4 permits comparison of the results from columns 2a and 2b of Table 2 with those for the same models re-run on the samples without India and without the five largest high-son-preference national populations. The bolded coefficients are ones for which the inferences differed across the samples. These discrepancies in the inferences support our hypothesis that the effects of fertility decline on gender gaps in children's well-being 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 samples that excluded high-son-preference populations. In the latter samples, a decline of one birth in total fertility was associated with increases of 0.06–0.08 in girls' mean *haz* scores at 0–36 months as well as reductions of 1.4–1.7 percentage points in the prevalence of stunting at these ages. In other words, in lower-son-preference populations, girls accrued greater nutritional benefits from declines in total fertility. Second, whereas a decline in total fertility 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 for boys), a decline in total fertility was not significantly associated with changes in the gender gap in early-childhood mortality in the samples that excluded high-son-preference populations (although most coefficients remained the same in sign). These findings corroborate the idea that an intensification effect of fertility decline on girls' excess mortality at ages 1–4 may be most pronounced in settings with an initially high son preference. Third, in the full sample, increases in women's median AFB were significantly associated with a *decline* in girls' initial excess mortality at ages 1–4 and an *increase* in girls' initial nutritional advantage (e.g., girls experienced faster declines in mortality and stunting than did boys); and, in the samples excluding high-son-preference populations, women's increasing median AFB remained associated with declines in the gender gap in both mortality at ages 1–4 as well as all but one measure of nutritional status. These findings support the idea that later first childbearing accrues greater reductions in mortality and stunting for girls than boys in high- and low-son-preference populations. Fourth, in the full sample, women's increasing median age at first birth was not associated with changes in the gender gap in vaccination coverage (initially favoring boys in 45 of 73 countries) but was associated with declines in this gender gap when high-son-preference populations were excluded. These findings suggest that removing the latter populations exposed a reversal in boys' advantage in vaccination coverage to favor girls in more gender-equitable populations and that later first childbearing may signal more substantial changes in gender norms in lower-son-preference populations.

Discussion

In this article, we have examined whether national changes in fertility regimes (from high total fertility and early median ages at first birth to lower total fertility and later median ages at first birth) are associated with: (1) aggregate increases in girls'

Table 4 Populated-weighted random-effect estimates of total fertility rate (TFR) and women's median age at first birth (AFB) on girls' well-being and gender gaps in children's well-being, full sample and sample without India and without the five most populous high-son-preference countries, Demographic and Health Surveys for 1985–2008

	TFR, Women Aged 15–49 at 0–4 Yrs. Before Survey									Median AFB, Women Aged 25–49								
	All Countries (from Table 2, column 2a)[a]			India Removed[b]			5 Most Populous High-Son-Preference Countries Removed[b]			All Countries (from Table 2, column 2b)[a]			India Removed[b]			5 Most Populous High-Son-Preference Countries Removed[b]		
	B	(SE)	*p*	B	(SE)	*p*	B	(SE)	*p*	B	(SE)	*p*	B	(SE)	*p*	B	(SE)	*p*
Girls' Well-being																		
Deaths to girls 1–4 yrs. per 1,000 1–4 yrs.	17.59	(1.78)	***	19.45	(1.71)	***	12.91	(1.77)	***	–10.43	(1.95)	***	–11.34	(1.92)	***	–7.44	(1.80)	***
% 12–23 months with specified vaccines	–8.31	(1.71)	***	–10.17	(1.59)	***	–7.42	(1.70)	***	5.80	(1.46)	***	5.44	(1.43)	***	4.56	(1.33)	***
Mean *haz* score <3 yrs.	**0.00**	**(0.03)**		**–0.06**	**(0.03)**	*	**–0.08**	**(0.03)**	*	0.12	(0.03)	***	0.08	(0.02)	***	0.12	(0.02)	***
% <3 yrs. stunted	**–0.40**	**(0.74)**		**1.41**	**(0.68)**	*	**1.71**	**(0.72)**	*	–2.09	(0.64)	**	–1.46	(0.57)	**	–2.54	(0.53)	***
Mean *waz* score <3 yrs.	0.03	(0.04)		–0.06	(0.04)		0.01	(0.03)		0.20	(0.03)	***	0.15	(0.03)	***	0.15	(0.02)	***
% <3 yrs. underweight	–1.08	(0.91)		1.13	(0.77)		–0.40	(0.83)		–3.77	(0.72)	***	–3.04	(0.58)	***	–3.48	(0.56)	***
Gender Gaps in Children's Well-being																		
Gap (girls–boys 1–4 yrs.) in risk of mortality	**–1.21**	**(0.60)**	*	**–0.46**	**(0.53)**		**–0.14**	**(0.55)**		**–1.17**	**(0.52)**	*	**–0.60**	**(0.45)**		–1.09	(0.40)	**
Gap (boys–girls 12–23 mo.) in % with specified vaccines	–0.23	(0.32)		–0.12	(0.31)		–0.08	(0.33)		**0.08**	**(0.29)**		0.04	(0.26)		**–0.57**	**(0.25)**	*
Gap (boys–girls <3 yrs.) in mean *haz* scores	–0.05	(0.01)	***	–0.03	(0.01)	*	–0.03	(0.01)	**	–0.04	(0.01)	**	–0.03	(0.01)	**	–0.02	(0.01)	*
Gap (girls–boys <3 yrs.) in % stunted	–1.61	(0.24)	***	–1.06	(0.21)	***	–1.26	(0.23)	***	**–0.60**	**(0.27)**	*	**–0.35**	**(0.20)**		**–0.21**	**(0.20)**	
Gap (boys–girls <3 yrs.) in mean *waz* scores	–0.05	(0.01)	***	–0.03	(0.01)	**	–0.03	(0.01)	***	–0.04	(0.01)	***	–0.04	(0.01)	***	–0.03	(0.01)	***
Gap (girls–boys <3 yrs.) in % underweight	–1.31	(0.19)	***	–0.92	(0.18)	***	–0.87	(0.17)	***	–0.46	(0.23)	*	–0.50	(0.17)	**	–0.36	(0.15)	*

Note: Refer to the footnotes to Table 1 for detailed definitions of the variables.

[a] All models are adjusted for country, time, and all socioeconomic conditions listed in the footnote to Table 2.

[b] The modeling strategy is identical to that used with "all countries," but observations from the indicated countries were excluded.

$*p \leq .05$; $**p \leq .01$; $***p \leq .001$

well-being and (2) changes in gender gaps in children's well-being, given the preexisting gender-normative environment.

Our results corroborate H1, showing that both declining fertility and women's later first childbearing were associated with declines in girls' mortality at ages 1–4 and increases in their vaccination coverage at ages 12–23 months. In the full sample, women's higher median age at first birth was associated systematically with improvements in girls' nutrition at ages 0–36 months, but changes in girls' nutrition per unit change in the TFR were not significant. Yet, in samples excluding high-son-preference populations, improvements in girls' nutritional status associated with declines in the TFR were significant. Thus, girls may benefit more from declines in total fertility only in populations with historically lower son preference.

Reductions in total fertility also predicted an *intensification* of gender gaps in mortality and nutrition (H2), whereas increases in women's median age at first birth were associated with more favorable outcomes for girls relative to boys (H3). Given that initial fertility decline is often driven by fertility restriction at older ages and that increases in women's median age at first birth occur later in the fertility transition (e.g., Knodel 1977; Sneeringer 2009), we interpret these findings as corroborating the theoretical pattern of association between changing fertility and changing gender gaps in children's well-being depicted in Scenario (d) of Fig. 3. That is, as total fertility initially declined, women who limited their fertility at older ages had more resources to invest in their children. Both the *increasing* advantage among boys in early-childhood mortality (where boys more often were advantaged initially; Table 3) and the *reduction* or *reversal* of girls' advantage in measures of nutrition (where girls more often were advantaged initially; Table 3) suggest that these newly available resources were at first disproportionately invested in boys, resulting in an intensification effect of initial fertility decline (corroborating H2). These greater investments in boys may have occurred more in the private than the public sphere, given that girls and boys experienced similar increases in vaccination coverage at ages 12–23 months as total fertility declined in all populations (Tables 2 and 4).

Subsequently, the postponement of fertility associated with an increase in women's median age at first birth may have signaled a broader shift in gender norms (e.g., Andersson et al. 2009; McDonald 2000; Presser 1971; Rosero-Bixby et al. 2009). As the value of girls (and women) rose, the relative investments in boys and girls also shifted. Consistent with H3, girls were able to catch up with boys (in the case of gender gaps in age 1–4 mortality) or to regain their natural biological advantage (in the case of nutritional statuses at ages 0–36 months).

The results of analyses based on samples excluding high-son-preference populations permit further interpretation of our findings with respect to H2 and H3. Corroborating H2, any intensification effects of declining total fertility on gender gaps in mortality appear to have been weaker in lower-son-preference populations (Table 4). Consistent with H3, any catch-up of girls relative to boys in vaccination coverage that was associated with women's later first childbearing appears to have been greater in lower-son-preference populations. Yet, girls' gains relative to those of boys in mortality and nutritional outcomes that were associated with later first childbearing appear to be similar across all settings. Overall, these findings suggest that the associations of fertility decline and women's later first childbearing with gender gaps in children's well-being may depend on preexisting levels of son preference, which arise from a range of non-fertility-related historical structural conditions. Still, the conditional influence of other (unobserved) characteristics that distinguished high-son-preference and low-son-preference populations cannot be ruled out.

Our analysis has several strengths. It is the only panel analysis using the DHS to assess how fertility declines are associated with diverse aspects of girls' well-being and gender gaps in children's well-being. We leveraged data from as many as 187 DHS from as many as 75 countries and spanning 23 years, linking these data to other national sources for a similar period. Finally, our analyses included rigorous and systematic adjustment for fixed- and time-varying sources of confounding in the relationships of interest. These strengths extend prior cross-national studies of the influences of fertility decline on girls' well-being, which have examined fewer outcomes for a narrower geographic scope or more limited time period.

Some aspects of our analysis suggest promising avenues for further research. First, because of our period of observation (1985–2008), we did not observe the full trajectory of changes in well-being that may arise from changes in fertility. As the DHS is extended, researchers may conduct longer time-series analyses on these and other aspects of well-being (e.g., women's average body mass index (BMI), extent of anemia, and access to prenatal care). Second, this was a cross-national time-series study, and the results should be interpreted in the aggregate. A complementary multilevel analysis might examine how meso-level changes in fertility, desired fertility, and son preference may have affected parental investments in sons' and daughters' well-being. Third, because the DHS are conducted in a nonrandom sample of lower-income countries, inferences to other countries should be made with caution and similar analyses should be undertaken with other national surveys. Finally, our analysis lacked direct measures of changes in national norms pertaining to fertility and gender preferences. Researchers might apply our analytical approach to study changes in fertility and our outcomes within countries having sufficiently variable sex ratios at birth and reliable panel data for districts.

Our findings suggest that fertility decline is associated with improvements in girls' well-being and broad declines in girls' risks of mortality and malnutrition, especially where son preference has historically been weaker. Yet, the influences of changing fertility regimes on gender gaps in children's well-being may vary. On the one hand, fertility decline may accrue equal benefits for boys and girls with respect to some intermediate needs for well-being, such as vaccination coverage. The concurrent expansion of public health infrastructures may explain some of this relationship, and further research exploring this possibility is warranted. 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. Yet, later increases in women's median age at first birth, which may signal (especially in less gender-biased populations) shifts toward more equitable gender norms, are associated with greater gains for girls than boys in survival, nutrition, and vaccination coverage. Thus, the influence of fertility decline on gender gaps in children's well-being may (1) depend on the extent of son preference as fertility initially declines and (2) vary across the stages of fertility transition as gender norms also shift. To avert an intensification effect of initial fertility declines, family planning programs in early transition settings could promote gender equitable investments in children.

Acknowledgments The authors thank the International Center for Research on Women (ICRW) and the Hewlett Foundation for their support. We also warmly thank members and consultants of the Fertility and Empowerment Research Network of ICRW for helpful comments on prior versions of this article. We especially thank Drs. Keera Allendorf, Sajeda Amin, Anju Malhotra, Karen Mason, and Amy Tsui. We also thank Ms. Emily Misch and Ms. Kimi Sato for their assistance with data preparation, and Ms. Francine Pope for assistance with manuscript preparation. This article was drafted while Drs. Nafisa Halim and Sarah Zureick-Brown were post-doctoral fellows and Ms. Kayla LaVilla was a Masters of Public Health student in the Hubert Department of Global Health, Rollins School of Public Health at Emory University. Any remaining errors are the responsibility of the authors.

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