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Selective Versus Generalized Gender Bias in Childhood Health and Nutrition: Evidence from India

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Abstract

We use data from India's National Family Health Survey (conducted in 1992–93, 1998–99, and 2005–06) to study gender discrimination across India among children aged 0–35 months. We focus on four measures of parental investment: (1) immunization, (2) received medical treatment for acute respiratory infection (ARI), (3) breastfed beyond 17 months, and (4) severe stunting. We contrast *generalized discrimination* that affects all daughters (vs. sons) with *selective discrimination* that posits enhanced discrimination among daughters who have sisters. We test which form of discrimination dominates and whether their effects are reinforcing. The main finding from this study is consistent and straightforward: daughters face discrimination relative to sons regardless of sibling composition and at both low and high birth orders. Further, net of general discrimination against daughters, we find little evidence of selective gender discrimination by birth order or sibling composition.

In 1987, Das Gupta published an influential article that introduced the concept of *selective discrimination*, i.e., gender discrimination that varies by the birth order and sex of a child's older siblings. Das Gupta showed that daughters with at least one sister had a higher risk of mortality, and she argued that the main pathway to excess mortality for second and third daughters was the inadequate parental allocation of key resources (such as nutrition and medical care) to those daughters. These daughters received fewer key investments from their parents because of the “deliberate failure to provide crucial inputs for sustaining life” (Das Gupta 1987, p. 86). In short, Das Gupta argued that selective discrimination results from conscious calculations and decisions about the relative value of each daughter.

In contrast, *generalized discrimination* affects all daughters and is rooted in the patriarchal social structure. Such discrimination is based on schemas in the “brain and in the world” (Johnson-Hanks et al. 2011) that specify that parents should provide superior care to sons relative to all daughters. Thus, all daughters face a form of “naturalized deprivation” wherein family members accept the gendered reasoning that deem daughters as “lesser beings” (Croll 2000).

Based on her descriptive analyses of a small sample from one district in Punjab, India, Das Gupta showed that selective, rather than generalized discrimination was likely at work. In this rural setting, she reported that women greatly valued having sons and did not want more than one daughter. A few subsequent articles have examined key mechanisms of selective discrimination (Muhuri and Preston 1991; Pande 2003; Mishra, Roy, and Retherford 2004; Jayachandran and Kuziemko 2011), while hundreds of articles have been written about son preference with its implied generalized discrimination against daughters.

Our article is the first to test formally which type of discrimination (selective or generalized) exists/dominates, or whether the two coexist and reinforce each other. While most of the literature on selective discrimination is based on rural samples from India (for exceptions, see Mishra, Roy, and Retherford 2004; Jayachandran and Kuziemko 2011), we use large nationally representative surveys from all of India and test for discrimination using a number of childhood health and nutrition measures for children aged 0–3 years. We also conduct the same analyses using the rural samples of India's National Family Health Survey (NFHS), and our results are essentially the same (available upon request). We examine data for a longer time period (1992–2005) than any other study, an important feature since the answer to the question of selective vs. generalized discrimination may vary over time. During our study period, contraceptive use and women's educational level increased greatly, while fertility and mortality declined substantially, with concurrent improvement and expansion of the public health infrastructure.

The case for generalized gender discrimination

Discrimination against girls and son preference have been enduring features of South Asian culture, embedded in and reinforcing patriarchy. Traditionally in India, a daughter's status within families has been low due to a combination of social, economic, and religious factors (Dyson and Moore 1983; Arnold, Choe, and Roy 1998). Underlying the strong traditions of patrilineal descent, patrilocal residence, and gender-biased inheritance laws is the age-old preferential treatment of males to preserve the family lineage (Dyson and Moore 1983; Jayachandran and Pande 2015; Ahlawat 2013). Sons cared for parents in old age, an offshoot of the custom of patrilocal residence. In line with the tradition of patrilineal descent, inheritance laws exclude or discriminate against daughters (Dyson and Moore 1983). Additionally, according to Hindu tradition, only sons can perform periodic religious rites for deceased parents so that their souls can attain salvation (Arnold, Choe, and Roy 1998). Sons are also valued for their greater economic utility: in an agricultural economy, sons would contribute more than daughters to the family economy through their labor. Sons can also command dowry and regular gifts from the wife's family. Given the multifaceted reasons for the strong parental preference for sons, differential treatment of sons and daughters is an expected, normalized feature of Indian society. In their quest for sons, parents may provide inferior resources to all daughters in the family, enabling them to provide disproportionate resources and advantages to sons.

To understand the generalized pattern of discrimination against daughters, it is informative to turn to localized studies. Even in rural areas in Tamil Nadu and Karnataka, two southern states that have completed the fertility transition and have weaker son preference than in

North India, women are averse to having daughters (Diamond-Smith, Luke, and McGarvey 2008; Sekher and Hatti 2010). To be sure, women in these studies emphasized that daughters provide emotional benefits to aging parents and are better caregivers than sons. Nonetheless, a significant proportion did not want any daughters because they were a financial drain on family resources due to the practice of dowry. Other ethnographic accounts narrate the anticipation for sons early in the reproductive career of wives, so that young wives can attain higher status within the family and couples can have smaller families than in the past (Croll 2000). These narratives document the muted response to a daughter's birth and the contrasting excitement surrounding the birth of a son (Croll 2000).

As Croll (2000) notes, parents perceive tangible differences between sons and daughters. Therefore, not only are daughters not substitutable for sons, but discriminating against them is anchored in parents' schemas about responsible care for their children. Jeffery and Jeffery (1996) argued that boys are believed to be more fragile starting from birth and in need of greater parental care to ensure their survival and good health. Accordingly, differential care in nutrition and health care has been extensively studied. Several studies report that daughters sit down to eat after the men of the house have eaten (Sekher and Hatti 2010), are consistently fed food that is considered less nutritious (Das Gupta 1987; Miller 1997), and are not given second servings (Miller 1997). Studies have shown that parents are less willing to spend money or provide high-quality care for daughters' health care expenses relative to sons (Basu 1989; Ganatra and Hirve 1994). Providing adequate health care such as routine immunizations and medical care for sickness is not merely related to the availability of such public health services. Indeed, immunization rates have markedly increased in recent years as the government provides free immunization through a well-established program. Rather, providing regular and timely immunization and medical care to every child requires a commitment of time and resources from parents. To the extent that parents subscribe to norms that devalue daughters relative to sons, they will be less inclined to expend those resources on daughters' health care, regardless of her siblings' sex.

Discriminatory treatment against daughters is not restricted to health care and nutrition. Studies show that sons are more likely to go to school and continue studying in higher grades than daughters (Sekher and Hatti 2010). When daughters are sent to school, they are sent to free, public schools, whereas sons are more likely to attend higher-quality private schools (*ibid.*).

The case for selective discrimination

Although there is a large literature on son preference in South Asia, few examine selective discrimination—i.e., discrimination conditional on a girl's birth order and sibling sex composition. In an early analysis of postnatal mortality in rural Uttar Pradesh, Simmons et al. (1982) described the selective allocation of resources within families to "assure the survival of some children at the expense of others" (p. 372) as a rational choice of parents. They noted that in communities with strong son preference, high fertility, and low prevalence of contraception, couples use mortality as a strategy to adjust their family size and the sex composition of their children. The circumstances of these societies make regulation of fertility more difficult than regulation of the number of surviving children and

their sex composition. Thus, in societies with a strong preference for sons, parents could practice female infanticide, selectively neglect unwanted daughters, or, in more recent periods, use sex-selective abortion.

A fundamental premise of selective discrimination is that daughters with at least one surviving sister receive more intense discriminatory treatment than those without a sister (Das Gupta 1987). Although sons with two or more brothers may also receive fewer health and nutrition inputs from their parents, the magnitude of these effects are smaller compared to those affecting daughters with multiple sisters (Pande 2003; Muhuri and Preston 1991).

Das Gupta (1987) showed that mortality rates among infants and children were substantially higher among daughters in higher birth orders and for those who had at least one surviving sister. She proposed that an important pathway to the higher mortality of certain daughters was the inadequate provision of key resources needed for survival in the early years: less nutritious food, poorer medical care, and lower expenditures on clothing (an important indicator of care given that this site experienced extremes in weather). Along the same lines, Muhuri and Preston (1991) used data from the Matlab research area in Bangladesh to study gender differences in mortality among children. They found that daughters with at least one older sister had an increased risk of dying compared to daughters with no siblings or with only older brothers.

Subsequent studies examined the pathways to daughters' excess mortality within the household. Pande (2003), using the rural sample of the 1992 NFHS, examined selective discrimination in immunization and stunting. Her research showed that redundant daughters (daughters born after one or more previous daughters) experienced poor outcomes. Compared to sons with two or more sisters, daughters with two or more sisters were more likely to be severely stunted and less likely to receive all immunizations. Mishra, Roy, and Retherford (2004) used NFHS 1992 and 1998 to demonstrate that gender bias in health care and nutritional status was greatest among daughters with no brothers. While they found inconsistent patterns of discrimination in nutrition outcomes, sons without brothers were much more likely to receive medical treatment for respiratory infections than daughters without brothers. Similarly, sons were less likely to be severely stunted when they had no surviving brothers, relative to daughters without brothers. In short, they argued that discrimination against girls was not generalized, but was based on their birth order and the sex composition of surviving siblings. More recently, using all three waves of the NFHS, Jayachandran and Kuziemko (2011) found that Indian mothers use the contraceptive effects of breastfeeding to limit or continue having children. At every birth order, mothers weaned their daughters early when they had not met their desired number of sons. On the other hand, when mothers have met their target family size and composition, they were more likely to continue breastfeeding. Our study also uses the NFHS and can be viewed as a re-analysis and an extension of these three studies.

Related arguments highlight the role of differential fertility levels by sex of previous children. As Bongaarts (2013) notes, parity progression in regions with strong son preference depends on the sex composition of current children. He further states that unbalanced sex ratios at birth provide "unambiguous evidence" of sex-selective abortion in

many Indian states. Moreover, evidence shows that the sex ratio at last birth, a particularly sensitive indicator of son preference, is highly elevated in most Indian states (ibid.). When sex ratios at birth are sensitive to the sex composition of existing sibships, we know that selective strategies are being used.

In addition to sex-selective abortions, Indian couples rely on differential stopping in their pursuit of sons. Daughters resulting from this strategy are likely at high risk, and several aspects of differential allocation of resources to higher-order daughters are salient for selective discrimination. First, the sex of the first child continues to behave as a random event, even though the 2005 NFHS provides some evidence that this pattern may be changing (Arnold, Kishor, and Roy 2002; Retherford and Roy 2003; Jha et al. 2011; Roy and Chattopadhyay 2012). At the same time, because Hindu parents gain goodwill by giving away a daughter in marriage, parents prefer at least one daughter (Arnold, Choe, and Roy 1998) and might not consciously discriminate against a first-born daughter. It is at higher parities that parents might make calculated prenatal and postnatal decisions such as whether to carry a girl fetus to term and to provide key inputs necessary for the daughter's survival and development. Second, given a strong desire for one or two sons, girls on average are born into larger families than boys (Barcellos, Carvalho, and Lleras-Muney 2014). Higher-order daughters, therefore, might receive fewer resources even if parents do not actively discriminate against them (Basu and De Jong 2010). Third, parents who already have one or more sons prefer to invest their resources in their son(s) rather than in higher-order daughters (Das Gupta 1987; Pande 2003; Mishra, Roy, and Retherford 2004). Fourth, higher-order daughters who have only sisters represent an "income shock" on parental resources (Jayachandran and Pande 2015); parents would rather conserve these resources for a son born in the future. For all of these reasons, girls born in higher orders and with sisters can expect to be targeted for discrimination by their parents.

Three hypotheses

Strong substantive arguments support both generalized and selective discrimination, and empirical evidence for both has been cited above. This leads us to propose three hypotheses and to conduct formal tests of whether the pattern of discrimination is generalized, selective, or both.

Hypothesis 1: Generalized discrimination: at a given parity, parents discriminate against all daughters.

Hypothesis 2: Selective discrimination: parents selectively target redundant daughters for discrimination. Thus, at higher birth orders, daughters with sisters will face worse outcomes than daughters with brothers.

Hypothesis 3: Generalized and selective discrimination: All daughters face discrimination, but redundant daughters face a double penalty: the generalized gender bias and selective bias because they have one or more sisters.

Data and methods

We use data from the National Family Health Survey conducted in 1992–93, 1998–99, and 2005–06 (referred to as 1992, 1998, and 2005 respectively). All three surveys are nationally representative samples of women of reproductive age and their households. In 1992, NFHS interviewed 89,777 ever-married women aged 13–49 living in 88,562 households.

Subsequently, the survey interviewed 89,199 ever-married women aged 15–49 from 91,196 households in 1998, and 124,385 women aged 15–49 from 109,041 households in 2005.

The surveys adopted a uniform sampling design in all study states, consisting of a multistage, systematic, stratified sample of households in both rural and urban areas. In rural areas, the sampling team first selected villages from the most recent census list. Primary sampling units (PSUs) of about 30 households were drawn systematically with probability proportional to size from each village (or set of adjacent villages when the village was too small). A similar approach was followed in urban areas. In the first stage, urban wards, cities, or towns were sampled with probability proportional to size. Next, from each ward, city, or town, a census enumeration block consisting of 150–200 households was selected. Finally, 15–60 households were selected within each block to constitute the PSUs in urban areas. Within each household, the survey teams interviewed eligible women who slept in the household the night before the survey. More details of the sampling design are available in the NFHS reports (IIPS 1995; IIPS and Macro 2000; IIPS and Macro 2007).

All surveys collected information on women's demographic and socioeconomic characteristics, including maternal and child health indicators. To assess childhood health and nutrition, the surveys also collected anthropometric and immunization reports for children under 4 years in 1992, under 3 years in 1998, and under 5 years in 2005. To make the results comparable across surveys, we restrict analyses to ever-married women aged 15–49 who were usual residents of the household and who had coresident children under age 3 (i.e., 35 months or younger). In addition, we restrict analysis to singleton births surviving at the time of the survey. Because the analytical period in each survey is the 35 months before the survey, there is a very modest potential for large sibships to contribute a greater share of observations. To elaborate, only 11 percent of the children in this sample are second siblings (number, and not birth order), and only 0.09 percent of children are third siblings. Further, for select analyses, we randomly chose one child per mother to validate our findings for a sample focused only on independent observations; these analyses produce results indistinguishable for those shown here.

Sibling composition is based on the index child's older surviving siblings as of the survey date. This operational decision makes our results conditional on any sex-selective abortions and on differential infant and child mortality. We address this issue in our concluding discussion.

Dependent variables

For every child born in the reference period, the surveys collected information from mothers on immunizations; episodes of acute illnesses such as upper respiratory infections, fever, and

diarrhea in the two weeks before the survey, and the treatment provided; breastfeeding; and height and weight measurements. We used this information to construct four outcome variables that are consistently measured in all three surveys: any immunization, received medical treatment for acute respiratory infection (ARI), prolonged breastfeeding, and severe stunting.

To qualify as fully immunized according to World Health Organization (WHO) guidelines for immunization, a child should receive the BCG vaccine at or soon after birth, three doses each of polio and DTP vaccines starting at 6 weeks and administered at 4-week intervals, and the measles vaccine at 9 months (WHO 2016). We constructed a binary variable to indicate whether a child aged 12–35 months received any immunization (any of the eight vaccine doses = 1) from immunization cards provided by their mothers, or from mothers' recall information when cards were not available. We chose children who were at least 1 year old at the time of the survey following WHO guidelines that all children should have these basic vaccinations by the time they are 1 year old. In analyses not shown here, we examined whether a child aged 12–35 months received all immunizations. The results of this analysis led to the same substantive conclusions as the analysis of receiving any immunization (results available upon request).

For medical treatment-seeking behavior, mothers were asked whether an index child suffered from ARI in the two weeks before the survey, and, if so, whether the child was taken to any public and private health facility for medical treatment. As the NFHS points out (IIPS and Macro 2000; IIPS and Macro 2007), a comparison of actual prevalence of ARI across survey years is not meaningful because the vast majority of ARIs are seasonal and the surveys were conducted at different times of the year. However, the analyses reported here would help identify the differences in health-seeking behavior between sons and daughters for acute infections. We constructed a binary variable to indicate whether the mother sought medical treatment for the index child aged 0–35 months when he or she was sick with ARI (sought medical treatment from private or public health facilities, not a pharmacy, shop, or traditional healer = 1).

The World Health Organization (WHO) and the Government of India currently recommend that children should be breastfed exclusively for up to 6 months, after which breastfeeding can be supplemented with appropriate nutrient-rich foods for up to 2 years or beyond (IIPS and Macro 2007; WHO 2016a). Breastfeeding is nearly universal in India, and the mean duration of breastfeeding is long. Yet, the duration of exclusive breastfeeding is very short (around 2–3 months) for both girls and boys. Only 69 percent of infants under 2 months, and only 28 percent of infants aged 4–5 months, were exclusively breastfed. Even in the first few days after delivery when it is important to establish breastmilk production, 57 percent of mothers gave their newborns something in addition to breastmilk to drink. Overall, girls are breastfed (not exclusively) for a slightly shorter duration than boys in all three surveys, although boys are not always provided age-appropriate solid foods (Mishra, Roy, and Retherford 2004). Because the median duration of breastfeeding is around 17 months, we constructed a dummy variable to indicate whether the mother breastfed the index child for more than 17 months, among children aged 18–35 months.

According to WHO, stunting is indicated by a low height for age (more than 2 standard deviations below the median for stunting, and more than 3 standard deviations below the median for severe stunting). Stunting reflects the long-term effects of undernutrition and is not easily reversible (WHO 2006). Close to half (48 percent) of children under 5 years in India are stunted, and India accounts for a third of all stunted children in the world (IFPRI 2016). We constructed severe stunting as a binary variable (severely stunted=1) among surviving children aged 7–35 months. In 1992, NFHS did not collect height data from five states surveyed in the first phase, referred to as Phase I states, due to lack of appropriate supplies to measure height at that time: Andhra Pradesh, Himachal Pradesh, Madhya Pradesh, Tamil Nadu, and West Bengal. Therefore, we analyzed severe stunting for the last two survey years: 1998 and 2005.

Independent variables

For each outcome variable, the focal independent variables are the combined categories of sibling composition and sex within a birth order. At every birth order (from 2 to 5), we include distinct combinations of the sex composition of surviving children (or sibling composition) and sex of the index child. In describing the categories of sibling composition within each birth order, we record the number and sex of siblings, but not the order in which they were born. For instance, for the third birth order, a girl born after a first-born son and a second-born daughter (bgG) is treated the same as a girl born after a first-born daughter and a second-born son (gbG), because in effect, as the index child being studied, she has one sister and one brother. Additional contrasts, including the order of siblings, were examined, but these contrasts did not provide substantively important or interpretable variation. Therefore, within birth orders, we present results with the sibling sex composition, but not the order in which older siblings were born. For models analyzing the first birth, the primary independent variable is simply the index child's sex.

Control variables include the child's age in months, mother's age at birth, mother's education (none, primary, secondary, and high school completed), whether she was employed and earned cash, father's education (none, primary, secondary, and high school completed), religion (Hindu, Muslim, and other), caste (scheduled caste or tribe, and other) whether she was exposed to regular media (weekly access to television and radio and monthly access to a movie), number of household co-residents, household wealth,¹ region (North, South, Central, East, and North East), and survey year.

Analytical strategy

We initially explored whether gender discrimination varied by survey year (1992, 1998, and 2005). Further, we tested for differential year effects in all models presented in tables below. Neither analysis revealed statistically significant or substantively plausible patterns of change. In addition, much of the literature on India cited above focuses on rural India. Thus, we have replicated all analyses below using only the NFHS rural sample, which closely

¹We used the constructed variables made available by DHS for the household wealth index and wealth quintiles. For details on the DHS strategy to code the wealth index, see IIPS and Macro International 2007; Rutstein et al. 1999; Rutstein et al. 2000.

parallel the results for the national sample presented in this paper. All of these supplemental analyses are available upon request. All estimates below are based on pooled data across surveys and reflect how combinations of sex and sibling composition are associated with differential outcomes for sons and daughters.

We begin by examining gender discrimination for children with no siblings, since it captures the essence of generalized discrimination: discrimination without any influence of sibling sex composition. Next, for each outcome and within each birth order (from 2 to 5), we estimate four logistic regression models to detect discrimination and identify its form. Table 1 provides the template for these models using the third birth order as an example. The row labels represent the sibling composition/sex of the index child. For instance, a third-born daughter (G) could have two sisters (ggG), one sister (bgG), or no sisters (bbG). Similarly, a third-born son could have two sisters (ggB), one sister (gbB), or no sisters (bbB), for a total of six combinations of sex and sibling composition. The first model, the “full model,” treats every sibling composition/sex contrast as unique. This baseline model allows for a test of whether all meaningful variation in these contrasts is captured by the more parsimonious alternatives. Model 1 takes the form:

$$\ln[p_{ij}/(1 - p_{ij})] = \beta_j + S_{ij}\alpha_1 + X_i\alpha_2 + \mu_{ij} \quad (1)$$

where p_{ij} is the underlying probability that child i in birth order j experiences a given outcome (for example, vaccination); β_j is the intercept; S_{ij} is a set of variables representing the birth-order-specific sex composition categories (for example, in the third birth order, the sex composition categories are as given above; α_1 is the effect of sibling sex composition on the probability of experiencing the outcome; X_i represents the set of controls, and μ_{ij} denotes the unobserved characteristics that affect the probability of experiencing the outcome.

The second model posits generalized discrimination (GD): discrimination against all daughters is equal (i.e., the meaningful contrast across all six categories is whether the index child is female vs. male, Hypothesis 1). Model 2 takes the form:

$$\ln[p_{ij}/(1 - p_{ij})] = \beta_j + GD_{ij}\alpha_1 + X_i\alpha_2 + \mu_{ij} \quad (2)$$

where GD is a binary variable indicating whether the child is female, and α_1 is the generalized effect of sex. The third model posits selective discrimination (SD): daughters with sisters face discrimination (Hypothesis 2), but first daughters do not. Model 3 can be expressed as:

$$\ln[p_{ij}/(1 - p_{ij})] = \beta_j + SD_{ij}\alpha_1 + X_i\alpha_2 + \mu_{ij} \quad (3)$$

where SD is a binary variable indicating whether the child is a girl *and* whether she has any sister(s), and α_1 is the effect of selective discrimination against girls with sisters. Finally, Model 4 allows for the additive effects of generalized and selective discrimination

(Hypothesis 3). That is, it posits that daughters with one or more sisters face a penalty above and beyond generalized discrimination. Model 4 is described as:

$$\ln[p_{ij}/(1 - p_{ij})] = \beta_j + GD_{ij}\alpha_1 + SD_{ij}\alpha_2 + X_i\alpha_3 + \mu_{ij} \quad (4)$$

All outcomes analyzed are binary, and estimated associations are presented as odds ratios. If the odds ratio is greater than 1 (by a factor of $\exp(m)$), exposure to that variable increases the odds of the outcome, relative to the reference (by a factor of $\exp(m)$). If the odds ratio is less than 1, the odds of the outcome are reduced by this factor. For each outcome and sibship/sex of the index child, we select a preferred model based on the joint considerations of the relative model fit, parsimony, and substantive plausibility. We follow the Bayesian model selection strategy proposed by Raftery (1995) that relies on the BIC statistic. BIC allows for comparisons between nested and non-nested models, and Raftery identifies rules of thumb for choosing one model over competitors depending on their relative BIC value (Raftery 1995, Table 7). The preferred model is indicated by the lowest BIC value, and the degree of evidence for the preferred model relative to an alternative is signaled by the difference in BIC values. Raftery identifies BIC differences that provide weak, positive, strong, and very strong evidence for choosing one model over a competitor. Given that the full model is a baseline representing the full set of contrasts, we expect the preferred model to be one of Models 2–4, three substantively plausible descriptions of the widely acknowledged gender discrimination in India.

Results

Table 2 shows the distribution of the outcome variables among children in the three surveys by sex. Nationally across the three years, a little over half (51–52 percent) of the children were male, and the sex ratio of boys to girls in this sample is 108. The prevalence of immunization increased substantially over time for girls and boys. Overall, 80.6 percent of girls and boys received at least one immunization. The difference in the proportion of girls and boys who were immunized was greatest in 1992 (about 4 percentage points). This difference decreased substantially by 2005, when over 90 percent of girls and boys received at least one immunization. Gender differences in medical treatment were more pronounced. Among children who had an ARI, a significantly higher proportion of boys than girls received medical treatment in every year, although the gender difference declined in 2005. Although boys were breastfed for a longer duration than girls, the differences were small (less than one month), but gradually increased across years. A slightly greater proportion of boys than girls were severely stunted in both 1998 and 2005, and these differences were statistically significant. More than a quarter of children aged 7–35 months were severely stunted in 1998. In 2005, this proportion declined substantially. Online Appendix I includes the distribution of key background characteristics of the household, mother, and index child, including child's sex, birth order, and sibling composition.²

²Appendices are available at the supporting information tab at wileyonlinelibrary.com/journal/pdr.

Table 3 shows estimates of generalized gender discrimination (odds ratios: Girl/Boy) among first-born children. These estimates show that even first-born daughters are less likely relative to sons to receive these important health inputs from their parents. For three of the four outcomes, a first-born girl fares worse than her male counterpart. Relative to a first-born son, a first-born daughter has a lower likelihood of receiving any vaccination, getting medical treatment for ARI, and being breastfed for more than 17 months. With respect to the fourth outcome, severe stunting, we find a female advantage: first-born daughters are less likely to be severely stunted relative to boys. We will revisit this anomalous finding below.

We present our analyses for the other birth orders by describing in detail results for one outcome, receiving any vaccination (models for the other outcomes are available in online Appendices II–VI). In Tables 4–7, we present the odds ratios of receiving any immunization for girls versus boys at the second through fifth birth orders. In each table we compare the four models described above and in Table 1.

Table 4 displays odds ratios (G/B) of receiving any immunization for the second child. Model 1, the full model, allows for independent effects of each of the four combinations, allowing for differing effects for sons who have a sister vs. a brother. The BIC statistic is higher by a value greater than 10 for this model, providing very strong evidence that other models better capture the significant variation in the sibling composition/sex of the index child (Raftery 1995, Table 7). Specifically, in Model 2 (H1: GD) relative to sons, daughters (bG and gG) are estimated to have significantly lower odds of receiving any immunization. Model 3 (H2: SD) shows that daughters with a surviving sister (gG) have lower estimated odds of being immunized compared to all other contrasts combined (i.e., to sons, and to daughters with a surviving brother). This model therefore shows evidence for discrimination targeted at the second daughter in the family. However, because Model 2 shows that even daughters with a brother face discrimination, SD (Model 3) misses a vital piece of information. The much lower BIC value of Model 2 (compared to Model 3) provides very strong evidence that the former should be preferred.

The fourth model is attractive substantively since it incorporates both forms of discrimination that have received prior empirical and substantive support (H3: GD + SD). We display estimates for both GD and SD factors, but obtaining their combined effect is straightforward. Because odds ratios are multiplicative,³ the total effects of GD+SD are .81 for all daughters and .76 for daughters with sisters (.81*.94). However, the BIC difference (Model 3 – Model 2 = 9.0) provides strong evidence that Model 2, generalized discrimination, should be preferred.

In sum, Table 4 provides unambiguous support for H1: generalized discrimination (Model 2) provides the preferred description of the sibling composition/sex effects.

Does this simple and unambiguous result hold at higher parities? Table 5 displays the OR (G/B) of any vaccination for a child with two siblings by sibling composition/sex. Here, Model 2 (GD) shows that relative to sons, all daughters (ggG, bgG, bbG) have lower odds of

³While log odds estimated from a logit model are additive, odds are multiplicative when exponentiated.

immunization (similar to the effect estimated in Table 4). Model 3 (SD) shows that daughters with at least one sister (ggG and bgG) are less likely to receive vaccination. Model 4 allows for the additive effects of GD and SD; the combined OR for ggG and bgG (.80*1.06) reflects GD and SD. The bbG contrast (.80) reflects only GD. As in the previous table, the comparison of BIC statistics provides strong evidence that generalized discrimination should be selected as the preferred model.⁴ The empirical result is simple and unambiguous: again generalized discrimination is the preferred model.

For children in the fourth birth order, Table 6 continues to show a discriminatory pattern in vaccination against all daughters (GD). Estimates from Model 2 (GD) constrain all daughters to have reduced odds of immunization relative to sons, an effect that is slightly stronger than the OR for the targeted daughters (SD) hypothesis. Again, Model 3 can be rejected in favor of Model 2 based on its lower BIC value (lower by 6.8). Likewise, Model 4 is rejected in favor of Model 2 (the BIC difference is 8.9). These BIC differences indicate that the evidence for preferring Model 2 is strong to very strong (see Raftery 1995, Table 7).

For children in the fifth birth order, Table 7 presents further evidence of generalized discrimination in being vaccinated. Model 2 (GD) shows that all daughters relative to sons have lower odds of immunization. Model 3 represents SD with two contrasts: for daughters with two or more sisters whose odds of vaccination are reduced by .73 compared to sons; and for daughters with only one sister, whose odds are reduced by .64 compared to sons. Again, Models 3 and 4 can be rejected in favor of Model 2 (GD) on the basis of a comparison of BIC statistics (a difference of 4.7 and 5.7, respectively) with moderately strong evidence (see Raftery 1995, Table 7).

To summarize, the empirical results in Tables 4–7 for vaccination are remarkably consistent. In each table, the preferred model is one of generalized discrimination (GD, Model 2). This model fits the data better than one positing only selective discrimination (SD, Model 3). Moreover, there is no empirical evidence that, net of generalized discrimination, selective discrimination plays a significant role (GD+SD, Model 4). Further, although the full model (Model 1) provides information about the variation in every sibling composition/sex category, such variation is parsimoniously captured in every case by Model 2. Therefore, more complicated characterizations of this variation (represented by the full model)⁵ are not warranted.

In Table 8, we ask whether the consistent generalized discrimination in immunization within every birth order is essentially equivalent. Panel A, column 1 shows these estimates for each birth order (equivalent to the ones estimated in Tables 3–7, Model 2). Panel B, column 1, constrains the effect to be pervasive across birth orders: daughters' likelihood of any vaccination is reduced by a factor of .88. The model with this equality constraint cannot be rejected on the basis of model fit comparisons (not shown). Thus, we conclude that the

⁴The BIC differences between Model 2 and its competitors (Models 1, 3, and 4) are, respectively 40.2, 6.8, and 8.7. [OK?] These differences with a sample size [OK?] of approximately 10,000 translate to very strong or strong evidence for preferring Model 2 (See Raftery 1995, Table 7). THIS IS FINE..

⁵This includes models that specify differential treatment among sons, e.g., that the first son would be most likely to receive key inputs (such as vaccination).

discrimination in immunization faced by all girls in every birth order is also pervasive across birth orders.

Is generalized discrimination also dominant and pervasive by birth order for other measures of child health and nutrition? Panel A in Table 8 (columns 2–4) presents generalized discrimination estimates by birth order for all indicators (the full sets of comparison models are in Appendices II–VI). We discuss these analyses in turn.

Column 2 in Panel A shows gender differences for medical treatment for ARI. These results parallel those for vaccination (in column 1, Panel A). Gender discrimination is clearly generalized, with no significant evidence of variation by parity. This result is also confirmed by model comparisons that parallel those above. Additionally, the preferred model for these data (see column 2, Panel B) indicates that daughters (compared to sons) are less likely by a factor of .77 to receive treatment for ARI (regardless of sibling composition or parity).

Column 3 in Table 8 shows these same estimates for extended breastfeeding. A model of generalized discrimination in Panel B produces estimates very similar to those in columns 1 and 2. Again we find no evidence that the strength of generalized discrimination varies by birth order. However, more detailed analyses (see Appendixes III and IV) do reveal associations at birth orders 3 and 4 that could be interpreted as selective discrimination. At these birth orders, a first daughter (i.e., one born after two or three sons) is breastfed longer than are daughters of the same birth orders who have sisters. We interpret this effect not as favorable treatment for a first daughter but as satisfaction with family size/sex composition. As Jayachandran and Kuziemko (2011) found using the same data we use, when Indian mothers reach their ideal family size and composition, they are more likely to continue breastfeeding. By many accounts, a family with two sons and a daughter is considered ideal (Dyson and Moore 1983; Arnold, Choe, and Roy 1998). Similarly, parents are ready to stop childbearing after three sons and a daughter. It is at these birth orders that parents are most inclined to stop childbearing after realizing their size and composition desires. Therefore, the first daughter after two or three sons is breastfed for a longer duration, not for the health benefits it provides her, but because these mothers are certain that they do not want more children (Jayachandran and Kuziemko 2011). Evidence that these daughters are not provided any superior treatment relative to other daughters can be seen in their lower odds of receiving vaccination and medical treatment for ARI. Thus, we argue that the estimates of generalized discrimination in column 3 of Table 8 best characterize the pattern of gender discrimination in extended breastfeeding.

Column 4 of Table 8 shows comparable estimates for severe stunting. But in this case we find a consistent advantage for daughters relative to sons. At each birth order in Panel A, daughters are less likely to be severely stunted than sons. As in columns 1–3, this effect can be constrained to be equal across all birth orders—daughters are less likely to be stunted by a factor of .87 (Panel B, column 4). How do we explain this theoretically anomalous result? We begin by noting that this result is not new (Gordon and Halileh 2013; Shroff et al. 2009; Subramanyam et al. 2010; IFPRI 2016; Mishra, Roy, and Retherford 2004; Mishra et al. 1999). Thus, while it is not empirically anomalous, it is inconsistent with our finding using the other three outcomes. A claim of pervasive discrimination against daughters, a

conclusion supported by most of the evidence presented here, must confront this result substantively and empirically.

We do so as follows. First, we propose that girls are more robust at birth and thus less likely to be severely stunted early in life. Many studies show that girls have a survival advantage in the first month of life (Wells 2000; Kruger and Nesse 2006; Katz et al. 2003). Others have shown the lack of gender differential in mortality in the first six months of life (Muhuri and Preston 1991). The gender disadvantage in stunting we expect among surviving children should accumulate over time due to both nutritional deprivation and heightened exposure to infections. Therefore, stunting is unlikely to be manifested in very young female infants, who are born robust. Rather, stunting would appear as girls age and suffer repeated health insults. Given this argument, the female advantage in survival shown in the early months of life shifts with age to a female disadvantage, one that is pronounced for children over 6 months of age (*ibid.*). To test this argument, in additional analyses (Figure 1 and Appendix VII) we examined whether the female advantage vis-à-vis severe stunting erodes and possibly reverses as the insults of disadvantage accumulate. Specifically, we tested for a sex-by-age interaction in the odds of severe stunting. This interaction is quite powerful (statistically and substantively): at the youngest ages examined (e.g., 7 months) the odds ratio (G/B) is .34 (indicating that the odds that a 7-month-old girl is severely stunted is reduced by a factor of .34 compared to a 7-month-old boy). This female advantage declines linearly with the child's age. By age 29 months the female advantage is no longer visible. This result is not sensitive to the functional form: modeling age as a linear or non-linear variable produces an identical substantive outcome. Thus even this apparently anomalous result can be reconciled as consistent with the overwhelming evidence presented here of pervasive, generalized discrimination against girls.

Summary and discussion

Selective child neglect for reasons such as determining family size and sex composition has occurred in many parts of the world for centuries (Scrimshaw 1978). Since the late 1980s, gender discrimination among young children in India was hypothesized to be conscious, selective, and targeted at certain daughters. Using more recent and nationally representative data from India, we tested whether gender discrimination in critical childhood health and nutrition outcomes was general or selective. The main finding from this study is consistent and straightforward: daughters face discrimination relative to sons regardless of sibling composition and at both low and high birth orders. Further, net of general discrimination against daughters, we find little evidence of selective gender discrimination by birth order or sibling composition.

Selective discrimination in nutrition and care in India is a critical problem because it could be directly linked to the excess mortality of daughters with sisters and those in higher birth orders. A proper test of selective discrimination requires that one control for general discrimination. As Table 1 shows, general and selective discrimination are intertwined in many sibling configurations. However, the first daughter can experience only generalized discrimination, since selective discrimination is directed at “redundant” daughters. Redundant daughters could be expected to face the cumulative effect of generalized and

selective discrimination. If one tests whether daughters with sisters face discrimination vis-à-vis all other children (H2 embodied in Model 3, Tables 4–7), then the answer is clearly yes. But if one controls on discrimination against all daughters (as in Model 4, Tables 4–7), then no additional effect of selective discrimination can be found. Only by explicitly and jointly testing for selective and generalized discrimination can we accurately gauge whether daughters with sisters are at greater risk than those without sisters. This surprising result can be aligned with some prior studies by the superior research strategy we use.⁶

We do not claim that selective discrimination never occurs. Such discrimination is clearly evident in unbalanced sex ratios (produced by sex-selective abortions) and in infant mortality in some contexts (Mosley and Chen 1984; Das Gupta 1987; Muhuri and Preston 1991). Our results can be aligned with this broader literature by positing that selective discrimination affects primarily the decision to use sex-selective abortion or to practice severe neglect/infanticide to achieve a desired sex composition. These extreme measures reflect the paramount importance of having a son or limiting the number of girls at particular stages of family formation. In contrast, daughters who are allowed to be born and to survive the early months of life are “wanted” (Short et al. 2001) or at least “tolerated” (Das Gupta and Bhat 1997). Given ingrained schemas that incorporate the differential value of sons and daughters, gender bias in caring for surviving children may be reduced, but not eliminated, by sex-selective abortion and neglect/infanticide (Short et al. 2001). This is evident in the pervasive (across birth orders), persistent (across time), but modest (in magnitude⁷) effect of generalized discrimination that we observe across a range of outcomes. In this context, daughters also adopt (consciously or unconsciously) the schemas that define their transient and burdensome presence in their natal homes. Caught in the interplay between the economic, social, and religious processes that mandate gendered roles, mothers and daughters alike “naturalize their deprivation” (Croll 2000). In this structural complex, they enact their inferior position in the family, reifying and recreating this patriarchal structure (Johnson-Hanks et al. 2011).

Do parents make strategic and conscious choices to discriminate against daughters? Or are they enacting pervasive and widely accepted schemas about the differential value and treatment of sons and daughters? We suggest that the answer is both.⁸ The former—

⁶Like the analyses in this article, Mishra, Roy, and Retherford (2004) stratify their analyses by birth order. For birth orders beyond the first, they use a variation of our full model, in which they include a composite variable composed of sex of the index child and his/her number of surviving brothers. As we demonstrated in our analyses, such a strategy would show the effect of selective discrimination, but without accounting for generalized discrimination. Pande (2003) analyzes selective gender discrimination by estimating main and interaction effects of sex, presence of brothers, and presence of sisters. Although her analysis shows a strong gender effect (what we refer to as GD) in models without interactions, it is vastly attenuated upon the inclusion of interaction terms. Particularly, girls with two sisters are much more likely to be severely stunted than are girls without siblings. In earlier analyses (not shown), we tried to replicate Pande’s findings. We were unable to obtain the same findings for severe stunting, but believe this could be due to the different analytic samples used in this article and by Pande. For instance, we use children aged 7–35 months, whereas she used children aged 6–47 months. She also used only the Phase 2 states for which height was measured in 1992. We used all states for which height was measured in 1998 and 2005. Our models, moreover, test interactive effects of sex and sibling composition both in full models and within and across every birth order. Further, we demonstrate that more parsimonious representations of these interactive models are possible and substantively meaningful.

⁷We seek to avoid the trap identified by Knodel and Jones (1996, p. 684): “By placing almost sole emphasis on gender inequality, demographers risk aligning themselves with a reactionary perspective that fails to emphasize the urgent need to remove obstacles to greater socioeconomic equity.” Indeed, while not the focus of this study, indicators of socioeconomic position (“control variables” in the current analysis) have associations with our focal outcome variables that are greater in magnitude than those we estimate for gender. Nevertheless, the gender gap in childhood outcomes is strong and pervasive.

conscious selective discrimination—operates primarily in selecting whether daughters are born and whether they survive the first few months of life. The latter—generalized discrimination via internalized schemas and the social context of daily life—operates in childhood among those daughters who are wanted/tolerated. Our research provides strong evidence of generalized discrimination among Indian children aged 0–35 months for vaccination, medical treatment for ARI, and breastfeeding beyond 17 months. Cumulative discrimination in these early months may even erode the significant health and survival advantage that girls have at birth. The broader literature provides evidence of selective discrimination depending on sibling composition, for example variation in sex ratios (suggesting sex-selective abortion) and infant mortality by sibling composition (Bongaarts 2013; Muhuri and Preston 1991).

Our findings and our attempt to integrate them into the existing literature identify three important questions for future research. First, in a particular context, whether India or elsewhere, can we show that selective discrimination operates prenatally and in the first six months of life, whereas generalized discrimination dominates subsequently? Second, India is a highly diverse country on multiple dimensions: its demographic profile varies by measures of rural/urban, socioeconomic status, religion, and region. Does the strength of generalized discrimination that we document vary by such factors? Much of the literature we cite focuses on India's rural population and our results hold there as well, but our analyses here pertain to all of India. Other differences remain unexamined. Claims of differential gender discrimination by sub-groups are common in the literature (Basu 1999; Borooah 2004; Dharmalingam and Morgan 2004; Mishra, Roy, and Retherford 2004). Our simple and straightforward characterization of generalized discrimination (as pervasive across birth order) lends itself to tests of variation across groups. Third, does the accumulation of disadvantage in key domains in the first three years of life dissipate the robustness of a girl's health and survival beyond the early years?

Despite significant social and economic development in India, childhood gender health inequalities remain embedded in patriarchy, a pervasive and enduring social structure. But patriarchy has additional pernicious effects on child health, even if equity in care of children could be obtained. For example, son preference can lead families to have additional children in the pursuit of sons. While not the focus of this article, our data clearly show that later-born children, both sons and daughters, are less likely to receive key resources compared to children of lower birth orders. In terms of the outcomes we study here, the disadvantage faced by a daughter (versus a son) is clearly important. But additional children born in the pursuit of sons increases the disadvantages faced by a child of either sex at a higher birth order. Thus, goals of improved child health overall, as well as greater gender equality in child health, would be among the benefits of reduced patriarchy.

Supplementary Material

Refer to Web version on PubMed Central for supplementary material.

⁸Selective and generalized discrimination are not necessarily produced by conscious choice and “naturalized” schemas, respectively. However, much of the previous literature has suggested these interpretations, and we find them reasonable.

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FIGURE 1. Odds ratio of severe stunting for girls relative to boys by age in months (NFHS 2, 3)

NOTES: Odds ratios estimated from logit regression model of severe stunting among children aged 7–35 months with sex of child, child's age, mother's age at birth, mother's education, exposure to media, mother's employment status, father's education, religion, caste, household wealth index, region, survey year, and birth order as controls. Estimates are obtained from interactions of sex of child with age.

TABLE 1

Model contrasts for different forms of gender discrimination, illustrated for third birth order

	(1)	(2)	(3)	(4)
Sibling composition/ sex of index child	Full	GD	SD	GD+SD
ggG	5	1	1	2
bgG	4	1	1	2
bbG	3	1	—	1
bbB	2	—	—	—
gbB	1	—	—	—
ggB	—	—	—	—
DF ^a	5	1	1	2

^a df: degrees of freedom

NOTE: Model description:

1. Full model: All contrasts unique
2. Generalized discrimination against daughters vs. sons
3. Selective discrimination against daughters with sisters vs. others
4. Generalized discrimination + Selective discrimination

TABLE 2
Sample sizes and percent distribution of children aged 0–35 months by outcome variables, NFHS 1, 2, and 3

	Pooled			1992			1998			2005		
	%	N		%	N		%	N		%	N	
Number of children												
Boys	51.9	43482		51.3	16605		52.4	13997		52.3	12880	
Girls	48.1	40225		48.7	15768		47.6	12720		47.7	11737	
Total	100.0	83707		100.0	32373		100.0	26717		100.0	24617	
Any immunization (12–35 months)												
Boys	81.9	24060		68.7	7560		86.0	8137		93.9	8363	
Girls	79.2	21384		64.6	6799		84.5	7238		92.9	7347	
Total	80.6**	45444		66.7**	14359		85.3*	15375		93.4*	15710	
Medical treatment for ARI (0–35 months)												
Boys	66.0	8539		68.4	3104		62.5	3213		68.0	2222	
Girls	59.8	6563		60.8	2390		56.4	2399		63.6	1774	
Total	63.2**	15102		64.9**	5494		59.8**	5612		65.9**	3996	
Duration of breastfeeding in months (18–35 months)												
Boys	69.9	14867		65.0	5154		73.4	4922		72.4	4791	
Girls	66.4	13011		61.5	4652		69.7	4233		69.3	4126	
Total	68.3**	27878		63.3**	9806		71.7**	9155		70.9**	8917	
Severe stunting (7–35 months)												
Boys	28.6	7398		—	—		27.7	3109		22.6	2098	
Girls	26.7	6376		—	—		26.2	2686		19.9	1661	
Total	27.7**	13774		—	—		27.0	5795		21.3**	3759	

* $p < 0.01$,

** $p < 0.001$; Traditional tests for differences by sex

NOTE: Weighted to account for clustered sample design.

TABLE 3

Effect of sex of index child on indicators of health and nutrition: First birth order (NFHS 1, 2, 3)

Sex of index child	Any vaccination	Medical treatment for ARI	Breastfed > 18 months	Severe stunting ^a
Odds ratios				
G (vs. B)	0.87 [*]	0.77 ^{***}	0.88 ^{**}	0.86 [*]
N	16216	7158	11704	10664

[†] $p < 0.10$,

^{*} $p < 0.05$,

^{**} $p < 0.01$,

^{***} $p < 0.001$

NOTE: All models control for child's age, mother's age at birth, mother's education, exposure to media, mother's employment status, father's education, religion, caste, household wealth index, region, and survey year.

^a Severe stunting was estimated from NFHS 2 and 3.

TABLE 4

Effect of Sibling Composition and Sex of Index Child on Receiving Any Vaccination: Second Birth Order (children 12–35 months old; NFHS 1, 2, 3)

Sibling composition/ sex of index child	(1)	(2)	(3)	(4)
	Full	GD	SD	GD+SD
	Odds ratios			
Gg	0.80 *	0.79 **	0.83 *	0.81 ** 0.94
Bg	0.87	0.79 **	—	0.81 ** —
Bb	1.03	—	—	— —
Gb	—	—	—	— —
DF ^a	3	1	1	2 2
BIC	9786.0	9761.8	9772.0	9770.8 9770.8
N	14568	14568	14568	14568 14568

* $p < 0.01$,

** $p < 0.001$

^adf: Degrees of freedom

Model description:

1. Full model: All contrasts unique
2. GD: Generalized discrimination against all daughters vs. sons
3. SD: Selective discrimination against daughters with sisters vs. others
4. GD+SD: Generalized discrimination + Selective discrimination

NOTE: All models control for child's age, mother's age at birth, mother's education, exposure to media, mother's employment status, father's education, religion, caste, household wealth index, region, and survey year.

TABLE 5
Effect of Sibling Composition and Sex of Index Child on Receiving Any Vaccination: Third Birth Order (children 12–35 months old; NFHS 1, 2, 3)

Sibling composition/ sex of index child	(1)	(2)	(3)	(4)
	Full	GD	SD	GD+SD
Odds ratios				
ggG	0.81 [*]	0.83[*]	0.89 [#]	0.80 ^{**} 1.06
bgG	0.93	0.83[*]	0.89 [#]	0.80 ^{**} 1.06
bbG	0.86	0.83[*]	—	0.80 ^{**}
bbB	1.00	—	—	—
gbB	1.00	—	—	—
ggB	—	—	—	—
DI ^{2a}	5	1	1	2 2
BIC	8115.4	8075.2	8082.0	8083.9 8083.9
N	9711	9711	9711	9711 9711

[#] $p < 0.10$,

^{*} $p < 0.05$,

^{**} $p < 0.01$,

^{***} $p < 0.001$

See text and notes to Table 4 for description of models.

TABLE 6

Effect of sibling composition and sex of index child on receiving any vaccination: Fourth birth order (children 12–35 month old; NFHS 1, 2, 3)

Sibling composition/ sex of index child	(1)	(2)	(3)	(4)
	Full	GD	SD	GD+SD
Odds ratios				
gggG	0.72 [*]	0.84 ^{**}	0.89 [#]	0.76 [*] 1.12
bggG	0.94	0.84 ^{**}	0.89 [#]	0.76 [*] 1.12
bbgG	0.73 ^{**}	0.84 ^{**}	0.89 [#]	0.76 [*] 1.12
bbbG	0.67 [*]	0.84 ^{**}	—	0.76 [*] —
gggB	0.79	—	—	—
bggB	0.79 [*]	—	—	—
bbgB	1.13	—	—	—
gggB	—	—	—	—
DF ^a	7	1	1	2 2
BIC	5860.8	5824.0	5828.0	5831.8 5831.8
N	5907	5907	5907	5907 5907

[#] $p < 0.10$,
^{*} $p < 0.05$,
^{**} $p < 0.01$,
^{***} $p < 0.001$

See text and notes to Table 4 for description of models.

TABLE 7

Effect of Sibling Composition and Sex of Index Child on Receiving Any Vaccination in the Fifth Birth Order (children 12–35 month old ; NFHS 1, 2, 3)

Sibling composition/ sex of index child	(1)	(2)	(3)	(4)
	Full	GD	SD	GD+SD
Odds ratios				
ggggG	0.96	0.77 ***	0.73**	0.99
bgggG	0.64**	0.77 ***	0.73**	0.99
bbggG	0.73*	0.77 ***	0.73**	0.99
bbbgG	0.63*	0.77 ***	0.64**	0.99
bbbbG	0.75	0.77 ***	—	0.99
bbbbB	0.95	—	—	—
gbbbB	0.86	—	—	—
ggbbB	0.79	—	—	—
gggbB	1.33	—	—	—
ggggB	—	—	—	—
DI ^{2a}	9	1	2	3
BIC	4124.6	4073.4	4078.1	4079.1
N	3779	3779	3779	3779

* $p < 0.05$,
** $p < 0.01$,
*** $p < 0.001$

See text and notes to Table 4 for description of models.

TABLE 8

Estimates of Generalized Daughter Discrimination (Odds Ratios, G/B) in Child Health and Nutrition, by Birth Order (NFHS 1, 2, 3)

Birth order	(1) Any vaccination	(2) Medical treatment for ARI	(3) Breastfed > 18 months	(4) Severe stunting ^a
Panel A: By birth order				
1	0.87 *	0.77 ***	0.88 **	0.86 *
2	0.79 ***	0.70 ***	0.80 **	0.90
3	0.83 *	0.82 **	0.80 ***	0.77 **
4	0.84 **	0.79 **	0.83 **	0.81 *
5	0.77 ***	0.78 *	0.70 ***	0.92
Panel B: Pooled birth orders				
Generalized discrimination (GD)	0.88 *	0.77 ***	0.88 **	0.87 *

[‡]
 $p < 0.10$,

*
 $p < 0.05$,

**
 $p < 0.01$,

 $p < 0.001$

NOTES: In Panel A, all models control for child's age, mother's age at birth, mother's education, exposure to media, father's education, employment status, religion, caste, household wealth index, region, and survey year. In Panel B, these same variables and their interactions with birth order are included as controls.

^aSevere stunting was estimated from NFHS 2 and 3.