# Associations Between Sex Composition of Older Siblings and Infant Mortality in India from 1992 to 2016

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A R T I C L E   I N F O

A B S T R A C T

*Background:* This study examines associations between sex composition of older siblings and infant mortality by sex, to guide efforts to address excess female infant mortality in India.
*Methods:* We conducted a retrospective cross-sectional study of infant mortality in India using four waves of data from the nationally-representative National Family Health Survey, collected between 1992 and 2016 (unweighted N = 338,504 for children aged 1—5). We used sex-stratified multivariable logistic regression models to assess the associations between sex composition of older siblings and risk of infant mortality.
*Findings:* Male infants with two living older sisters and no living older brothers had lower odds of infant mortality relative to those with one living older brother (e.g., 2015—16 AOR 0.62, 95% CI 0.50—0.76); this effect was significant for boys across all waves of data but was not seen for girls in any wave. Exploratory models focused on third order births found that boys were less likely than girls to die in infancy if born subsequent to two older sisters (2015—16 AOR 0.48, 95% CI 0.31—0.74); analysis of crude prevalence data indicated that this converts into a 64% greater risk for infant mortality for girls relative to boys in this third-order group.
*Interpretation:* Higher birth order males with older sisters have greater protection against infant mortality, a finding that has persisted for over 25 years. To address ongoing gender inequities in infant survival in India, greater focus is needed to support higher birth order girls and social norm movements against son preference.
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## 1. Introduction

Infant mortality in India has declined steadily over the past 25 years, dropping from 79 deaths per 1000 live births in 1993, to 41 deaths per 1000 births in 2016, yet concern remains over a higher than expected female mortality rate, and a female mortality ratio that has worsened over time [1,2]. Studies of sex differences in infant and child mortality in India have found that girls have a higher under-5 mortality relative to boys, and suggest this excess female mortality translates into between 216,300 to 329,900 excess deaths of girls under-5 in India each year [2–5]. Among infants, which comprise the majority of under-5 deaths, evidence of gender inequity is less clear. Analysis of nationally representative 2005—06 data and 2015—16 data from India found that boys were more likely than girls to die in the neonatal period

(0—1 month) [5,6], as would biologically be expected [7]. However, girls were more likely than boys to die in the postnatal period (1—12 months) in 2005—06, contrary to biologic expectation, though no sex difference in postnatal mortality was seen in 2015—16 [5,6]. These findings suggest the social norm of son preference, which contributes to excess female infant mortality, may be declining [5].

While these findings provide insight into targets for policy and practice interventions to help address excess female infant mortality in the country, they are insufficient to target those household contexts where such risk may be greatest. Differential effects have been shown by wealth, with recent data showing greater excess female postnatal mortality for poorer relative to richer segments of society [5]. Birth order and sibling composition may also affect sex differences in infant mortality in India, but has received less attention in the literature. Previous research from India documents differential associations between sex composition of siblings and child malnutrition and immunization by sex of the child [8–11]. Findings from these studies suggest that, among girls, having older brothers increases risk of wasting (acute malnutrition) [9] and having older sisters increases risk of stunting (chronic

## Research in context

*Evidence before this study*

Published research and national data document a higher than expected female infant mortality rate and a worsening female mortality ratio over time in India. Previous studies have also found associations between number and sex composition of older siblings and health outcomes including malnutrition and immunization. One previous study has examined older sibling number and sex composition and its association with infant mortality; it found that older sisters were associated with increased risk of female infant mortality, and found no sibling effects of male infant mortality. However, this study was based on decade-old data and did not include consideration of potentially confounding demographic and health systems factors.

*Added value of this study*

This study builds on previous research by examining the association between older sibling composition and infant mortality by relying on recent nationally representative data, using past one to five-year births, controlling for potentially confounding demographic and health factors, and examining trends in these associations over time.

*Implications of all the available evidence*

Our study finds that older sibling composition is associated with infant mortality, and that this association differs by sex, with greater protection from infant mortality for boys born subsequent to two sisters. Further, this finding appears to be strongest for the wealthiest populations. This study and other literature suggest a persistent trend of greater protection of male relative to female infants, maintained for the past two decades and strongest for third order births occurring subsequent to two older sisters. These findings indicate an ongoing social norm of son preference affecting differential investments in child survival by sex. Implications of these findings are that existing efforts to eliminate sex ratio imbalance and excess female mortality remain inadequate. Broader social norm change focused on eliminating son preference is needed, as is greater focus from clinical and community care efforts for higher birth order girls in India.

malnutrition) [9,10] and not receiving any or all required immunizations [10,11]; for boys, having older sisters was positively associated with immunization [10]. Eldest boys appear to be in the strongest position of parental investment, with the least risk of stunting [8]. Studies on sibling composition and infant mortality in India found that having older sisters increases risk of female infant mortality regardless of wealth; no sibling effects on male infant mortality were observed [6,12,13]. Last born children were also most vulnerable to infant mortality [6]. These studies were all limited to data from 10+ years ago, and the infant mortality focused study aggregated all births, impeding recent cohort considerations in analysis. More research is needed to understand these issues using more recent data, but also with consideration of whether we are seeing improvements over time.

This study extends prior research by examining the associations of sex composition of older siblings on infant mortality using sex-disaggregated data on the infant, to guide policies addressing child sex ratio imbalances and excess female infant mortality in India. First, we assessed the associations between sex composition of older siblings and infant mortality for girls and boys in the most recently available national data from India (2015—16), and then secondarily assessed whether engagement in essential maternal and newborn care

(antenatal care, skilled birth attendant, and breastfeeding) affect observed sibling effects. Then, we tested associations between sex composition of older siblings and infant mortality, by sex of infant, across four waves of data, from 1992—93 to 2015—16, to determine if observed associations have changed over time. These findings offer insight into whether son preference is declining in India, and also can guide health systems to target households with greatest risk for female infant mortality.

## 2. Methods

### 2.1. Data Source

We conducted a retrospective cross-sectional study of infant mortality in India utilizing National Family Health Survey (NFHS) data from 1992 to 1993 (NFHS-1) [14], 1998 to 1999 (NFHS-2) [15], 2005 to 2006 (NFHS-3) [16], and 2015 to 2016 (NFHS-4) [1]. Nationally representative household-based samples were created for each wave through a stratified, multistage sampling strategy, described in detail elsewhere [1]. All waves capture data on fertility, mortality, family planning, utilization of healthcare services, and other related indicators.

We limited the study sample to births within the one to five years prior to interview; more recent births were the focus to provide insight into practices and son preference in close proximity to the time of interview. We excluded past year births to allow for completion of the window of infant mortality (defined as death before first birthday) risk. For analyses involving prenatal care indicators, we limited our sample to most recent births in the one to five years, as questions regarding certain care indicators were asked only with reference to the most recent birth. We excluded multiple births (e.g. twin, triplet) and births to never-married mothers (excluded N = 12,057 births, 3.7% of total sample). The sample ultimately included 213,638 children born to 168,274 mothers for NFHS-4 data. The same criteria were applied to NFHS-1 (n = 49,041 children; n = 36,855 mothers), NFHS-2 (n = 46,400 children; n = 35,060 mothers), and NFHS-3 (n = 41,382 children; n = 32,697 mothers) data for cross-wave analyses.

### 2.2. Measures

The primary outcome was infant death within the first year of life. The primary independent variable was a five-level categorization of living older sibling composition at time of birth, defined as first born / no living older siblings, one living older brother only, one living older sister only, two living older brothers only, two living older sisters only, and a mix of older living brothers and sisters (including having one older brother and one older sister, and having any combination of three or more older siblings) [10].

Covariates were considered for inclusion based on two criteria: first, demonstrated association with infant mortality in previous research [1,17], and second, availability and low missingness across all four waves of NFHS data. All potential variables that met these criteria were included in our analyses, without further model building. We thus included maternal and household demographics, including mother's age at birth (linear and squared term, to account for non-linear relationship between age and infant mortality), age at marriage, education, religion, scheduled tribe/scheduled class designation, household structure (nuclear or non-nuclear), household asset-based wealth quintile, rural/urban area of residence, and national region of residence. Maternal education was defined using total years of completed education and categorized as no formal education, any primary education, some secondary education (not completed), and completed secondary education or more. Religion was categorized as Hindu, Muslim, Christian/Other/No religion. Household structure was defined as nuclear/non-nuclear; in in NFHS-3 and NFHS-4 the household structure was provided directly and for NFHS-1 and NFHS-2 nuclear households were defined based on household rosters as those containing

only head of households, spouse of head of household, and never-married children or adopted/foster children. We categorized regions as north, central, east, northeast, west, and south based on state, per NFHS-4 categorizations [1]. We also included two birth-level characteristics available from birth history items, birth spacing and an indicator of a prior sibling having died before birth, as covariates. Birth spacing was dichotomized as less than two years since previous birth or two or more years since previous birth/first born.

For analyses limited to most recent births in the past one to five years, we included available essential maternal and newborn care variables, including antenatal care (ANC), skilled birth attendance (SBA), and breastfeeding initiation timing, again based on their association with infant mortality in India [1,17]. Antenatal care was categorized as having received four or more antenatal care visits, one to three antenatal care visits, or no antenatal care visits. Skilled birth attendance was indicated if the mother reported that there was a doctor, nurse, Auxiliary Nurse Midwife (ANM), midwife, Lady Health Visitor (LHV), or "other healthcare professional" present at the birth. We dichotomized breastfeeding initiation as occurring within one hour of birth or not.

### 2.3. Data Analysis

We constructed a series of four sex-stratified analyses to examine sex-specific associations between sibling composition and infant mortality, and whether those associations changed over time. First, we used an adjusted logistic regression model of all births in the one to five years prior to interview for NFHS-4 to assess the association of sex composition of older siblings with infant mortality, stratified by sex of the infant. Second, we limited the sample to most recent births and included essential maternal and neonatal care covariates to examine the potential mediating effect of care-seeking behaviors on sibling order's association with infant mortality. Third, to examine trends over time, we constructed comparable sex-stratified models for the three previous NFHS waves. Fourth, we combined these four waves of data into a pooled model and included a wave-sibling composition interaction term to assess statistical significance for changes over time in the relationship between sibling composition and infant mortality, over the period of 1992—93 to 2015—16. These four analyses constituted our initial analytic plan developed a priori to address our research question; unadjusted models were also conducted for all variables included in the above models of NFHS-4 data as a result of peer review feedback.

We conducted two additional sets of exploratory analyses to further explore the relationship between sibling composition and mortality. We constructed a series of models stratified by each sibling composition group in NFHS-4 data on all births in the past one to five years. These five models pooled both sexes and included child's sex as a covariate to examine sex differences in infant mortality within the same older living sibling composition. We ran these models in two ways: including all children born in the previous one to five years and limited to the most recent child. We then conducted a second set of exploratory analyses assessing sex differences in infant mortality stratified by birth order. Two models were constructed: one for second born children, and one for third born children (fourth born or higher order children were excluded in this analysis). Sex composition of older siblings was included as a covariate, as was an interaction term between sex of infant and the sibling variable. Again, models were run in two ways: including all children born in the previous one to five years and limited to the most recent child.

Finally, after all above analyses were complete, additional post-hoc analyses were conducted to further explore significant findings of interest. These are described in detail below.

All analyses accounted for survey design and weighting, and were conducted using STATA 15.1. Ethical approval for NFHS-4 data collection was granted by Institutional Review Boards at the International Institute for Population Sciences and ICF. Exempt status for this analysis of publicly available, de-identified data was given by the University of California San Diego.

## 3. Results

The infant mortality rate (IMR) among the NFHS-4 study sample was 36.7/1000 live births (39.6/1000 live births among males and 33.5/1000 live births among females, demonstrating a steady and dramatic decline since NFHS-1 for both males (72.3/1000 live births) and females (68.3/1000 live births (Table 1). Among all births in the one to five years prior to NFHS-4, 41% were first born, 16% had 1 living older brother, 17% had one living older sister, 4% had two living older brothers, 8% had two living older sisters, and 15% had a mix of older living brothers and sisters. Sibling composition has also changed across study waves, reflecting a shift toward smaller family size. First-born children comprised a significantly larger portion of study children over time, while those with a mix of older living brothers and sisters (which includes those children who are the fourth birth order or higher) fell over time. In NFHS-4, sex differences were seen in IMR based on sibling composition (Appendix Table 1). While the overall female to male sex ratio IMR is 0.85 (i.e. 15% lower for females than males), the sex ratio IMR is 1.41 for third order births born after two brothers (i.e., 41% higher for females than males). The sex ratio IMRs for infants born after a male as a second order infant or subsequent to male and female children were approximately 1.00, again suggesting that girls are also at higher risk for female mortality for reasons beyond what would be expected biologically.

In the most recent round of data (NFHS-4, 2015—16), sibling order was significantly associated with infant mortality in both crude and adjusted models (Table 2; Appendix Table 2). For both males and females, having no living older siblings conveyed significantly greater risk of infant mortality than having one living older brother (male AOR 1.82, 95% CI 1.55—2.14; female AOR 1.86, 95% CI 1.58—2.20). While having two living older sisters and no living older brothers was significantly protective for male infants (AOR 0.62, 95% CI 0.50—0.76), for female infants, there was no significant association between being of second or higher birth order and infant mortality. Mother's age at birth, mother's education, household wealth, region, birth spacing, and having a sibling who died previously were also all significantly associated with infant mortality risk for both males and females.

These trends were maintained among the subsample of most recent births, even after adjusting for essential maternal and newborn care indicators (Table 3, Appendix Table 3). Having no living older siblings conveyed the greatest risk of infant mortality for both males and females (male AOR 1.46, 95% CI 1.13—1.89; female AOR 1.54, 95% CI 1.16—2.03), effect sizes slightly attenuated relative to that seen in the larger model and crude analysis. For males, having two living older sisters and no living older brothers remained significantly protective (AOR 0.59, 95% CI 0.43—0.82), with an effect size comparable to that seen in the primary model as well as the crude analysis. For females, there remained no significant association between sibling composition and infant mortality risk among those who were of a second birth order or higher. Neither having received ANC nor SBA were significantly associated with infant mortality risk, but breastfeeding initiation within 1 h of birth was strongly protective for both males and females (male AOR 0.26, 95% CI 0.22—0.30; female AOR 0.29, 95% CI 0.25—0.34).

Analyses across waves found similar relationships between sibling composition and infant mortality over time (See Fig. 1). There were no significant differences in association between sibling order and male infant mortality over time (i.e. no significant wave by sibling interaction) (Appendix Table 4). Among females, the relationship between sibling composition and infant mortality differed significantly over time in two instances: having two older living brothers (AOR 1.55, 95% CI 1.03—2.35) or a mix of older living brothers and sisters (AOR 1.44, 95% CI 1.10—1.89) conveyed significantly more risk relative to having one older living brother at NFHS-4 than at NFHS-1.

**Table 1**

Study population characteristics, India, 1992–1993, 1998–1999, 2005–2006, 2015–2016. Ns unweighted; percentages, means, and infant mortality rates survey weighted.[a]

| | NFHS-1 1992–1993 (N = 47,254) | NFHS-2 1998–1999 (N = 44,840) | NFHS-3 2005–2006 (N = 40,865) | NFHS-4 2015–2016 (N = 205,545) |
|---|---|---|---|---|
| **Infant mortality rate (Deaths per 1000 live births)** | | | | |
| All children | *70.4* | *63.2* | *53.4* | *36.7* |
| Male children only | *72.3* | *63.8* | *51.9* | *39.6* |
| Female children only | *68.3* | *62.6* | *55.0* | *33.5* |
| **Characteristics (N (%))** | | | | |
| **Sibling Composition** | | | | |
| No living older siblings | 14,584 (30.5) | 13,897 (30.7) | 13,982 (32.5) | 80,355 (40.8) |
| 1 living older brother | 6421 (13.3) | 6305 (14.3) | 5960 (14.3) | 31,864 (16.0) |
| 1 living older sister | 6156 (13.1) | 6143 (13.8) | 6026 (14.2) | 32,974 (16.6) |
| 2 living older brother, no older sisters | 2911 (6.2) | 2587 (5.8) | 1957 (4.9) | 8174 (3.6) |
| 2 living older sisters, no older brothers | 3605 (7.5) | 3502 (7.7) | 3182 (7.9) | 16,727 (7.8) |
| Mix of older living brothers & sisters | 13,577 (29.3) | 12,406 (27.7) | 9758 (26.1) | 35,451 (15.1) |
| **Mother characteristics** | | | | |
| Mother's age at birth *(mean)* | *23.2* | *22.9* | *23.4* | *23.8* |
| Mother's age at marriage *(mean)* | *16.5* | *16.7* | *16.3* | *18.4* |
| **Mother's education** | | | | |
| None | 28,491 (65.7) | 24,184 (57.7) | 17,037 (51.0) | 66,249 (31.0) |
| Any primary | 11,344 (21.9) | 12,328 (26.1) | 12,098 (27.6) | 66,860 (31.9) |
| Some secondary, not completed | 4730 (8.0) | 5186 (10.1) | 6768 (13.2) | 38,109 (18.9) |
| Completed secondary or more | 2523 (4.4) | 3122 (6.1) | 4962 (8.2) | 34,327 (18.2) |
| **Household characteristics** | | | | |
| **Household structure** | | | | |
| Nuclear | 20,143 (40.9) | 20,035 (43.6) | 20,329 (48.8) | 88,918 (42.1) |
| Non-nuclear | 27,111 (59.1) | 24,805 (56.4) | 20,536 (51.2) | 116,627 (57.9) |
| **Religion** | | | | |
| Hindu | 35,906 (79.5) | 33,212 (79.0) | 28,151 (78.1) | 148,237 (78.6) |
| Muslim | 6372 (15.4) | 6596 (15.9) | 6839 (17.2) | 32,509 (16.7) |
| Christian/ Other / No religion | 4976 (5.1) | 5032 (5.0) | 5875 (4.6) | 24,799 (4.8) |
| **Scheduled tribe/scheduled caste** | | | | |
| No | 34,985 (77.2) | 29,516 (70.0) | 26,640 (69.5) | 124,306 (67.9) |
| Yes | 12,269 (22.8) | 14,901 (30.0) | 13,872 (30.5) | 79,908 (32.1) |
| **Household wealth quintile** | | | | |
| Poorest | 8774 (22.5) | 8997 (24.3) | 7332 (25.6) | 54,750 (25.5) |
| Poorer | 8867 (22.0) | 8814 (22.4) | 7594 (22.5) | 48,584 (21.9) |
| Middle | 9651 (20.9) | 9288 (20.0) | 8480 (19.8) | 40,870 (19.7) |
| Richer | 10,778 (19.4) | 9699 (18.6) | 8933 (17.8) | 34,042 (18.2) |
| Richest | 9184 (15.2) | 8042 (14.7) | 8526 (14.3) | 27,299 (14.7) |
| **Region** | | | | |
| North | 10,591 (11.7) | 10,532 (12.8) | 7346 (12.9) | 38,550 (13.2) |
| Central | 11,204 (29.6) | 9887 (29.5) | 9271 (29.8) | 59,411 (29.6) |
| East | 7870 (22.3) | 7743 (22.0) | 6443 (25.2) | 42,783 (25.5) |
| Northeast | 5337 (4.5) | 6288 (3.7) | 7591 (3.7) | 29,806 (3.6) |
| West | 4991 (13.2) | 4397 (13.1) | 4456 (12.6) | 14,495 (12.8) |
| South | 7261 (18.7) | 5993 (18.8) | 5758 (15.8) | 20,500 (18.1) |
| **Urban residence** | | | | |
| Rural | 34,439 (77.3) | 33,297 (78.0) | 25,309 (74.4) | 156,574 (71.6) |
| Urban | 12,815 (22.7) | 11,543 (22.0) | 15,556 (25.6) | 48,971 (28.4) |
| **Sibling-related characteristics** | | | | |
| **Birth spacing** | | | | |
| Less than 2 years since previous birth | 9867 (20.6) | 9778 (21.8) | 7923 (20.2) | 35,310 (17.0) |
| 2 years or more since previous birth/1st born | 37,387 (79.4) | 35,062 (78.2) | 32,942 (79.8) | 170,235 (83.0) |
| **Older sibling died prior to birth** | | | | |
| No | 37,194 (76.5) | 36,288 (79.9) | 34,605 (81.9) | 185,924 (90.8) |
| Yes | 10,060 (23.5) | 8552 (20.1) | 6260 (18.1) | 19,621 (9.2) |
| **Healthcare factors: Available for most recent births only** | | | | |
| ***Most recent birth N*** | **28,678** | **27,926** | **26,647** | **139,463** |
| **Received antenatal care visits** | | | | |
| None | 8481 (36.7) | 5812 (34.4) | 4976 (23.2) | 25,182 (17.0) |
| 1 to 3 | 8452 (35.0) | 6192 (35.8) | 9218 (39.0) | 47,374 (31.1) |
| 4 or more | 7717 (28.4) | 5631 (29.8) | 12,154 (37.8) | 65,492 (51.9) |
| **Birth had skilled birth attendant** | | | | |
| No | 15,261 (64.9) | 10,051 (58.7) | 11,872 (51.6) | 28,187 (17.0) |
| Yes | 9324 (35.1) | 7639 (41.3) | 14,755 (48.4) | 111,194 (83.0) |
| **Breastfeeding initiated within 1 h** | | | | |
| No | 19,509 (86.5) | 11,917 (76.9) | 12,883 (62.0) | 34,169 (28.4) |
| Yes | 3968 (13.5) | 4478 (23.1) | 11,442 (38.0) | 92,179 (71.6) |

[a] Only non-missing values reported.

**Table 2**
Multivariate factors associated with infant mortality in all births in previous 1−5 years, India 2015−2016.

| | Males only (unweighted n = 103,990) | | Females only (unweighted n = 95,997) | |
|---|---|---|---|---|
| | Adjusted odds ratio | 95% confidence interval | Adjusted odds ratio | 95% confidence interval |
| **Sibling Composition** | | | | |
| No living older siblings | 1.82 | 1.55,2.14 | 1.86 | 1.58,2.20 |
| 1 living older brother | Ref | Ref | Ref | Ref |
| 1 living older sister | 0.89 | 0.75,1.05 | 1.14 | 0.95,1.38 |
| 2 living older brother, no older sisters | 1.03 | 0.80,1.32 | 1.10 | 0.85,1.44 |
| 2 living older sisters, no older brothers | 0.62 | 0.50,0.76 | 1.15 | 0.91,1.44 |
| Mix of older living brothers & sisters | 0.87 | 0.73,1.04 | 1.19 | 0.98,1.44 |
| **Mother characteristics** | | | | |
| Mother's age at birth | 0.84 | 0.79,0.89 | 0.86 | 0.80,0.92 |
| Squared mother's age at birth | 1.00 | 1.00,1.00 | 1.00 | 1.00,1.00 |
| Mother's age at marriage | 1.00 | 0.98,1.01 | 0.99 | 0.98,1.00 |
| Mother's education | | | | |
| None | Ref | Ref | | |
| Any primary | 0.88 | 0.79,0.98 | 1.03 | 0.92,1.15 |
| Some secondary, not completed | 0.75 | 0.64,0.87 | 0.78 | 0.66,0.92 |
| Completed secondary or more | 0.62 | 0.51,0.75 | 0.68 | 0.55,0.83 |
| **Household characteristics** | | | | |
| Household structure | | | | |
| Nuclear | Ref | Ref | Ref | Ref |
| Non-nuclear | 0.99 | 0.91,1.08 | 1.02 | 0.93,1.13 |
| Religion | | | | |
| Hindu | Ref | Ref | Ref | Ref |
| Muslim | 0.98 | 0.85,1.12 | 0.90 | 0.78,1.04 |
| Christian/ Other / No religion | 0.94 | 0.73,1.21 | 0.82 | 0.64,1.06 |
| Scheduled tribe/scheduled caste | | | | |
| No | Ref | Ref | Ref | Ref |
| Yes | 1.05 | 0.96,1.15 | 0.90 | 0.81,0.99 |
| Household wealth quintile | | | | |
| Poorest | Ref | Ref | Ref | Ref |
| Poorer | 0.95 | 0.85,1.06 | 0.88 | 0.78,1.00 |
| Middle | 0.89 | 0.77,1.02 | 0.95 | 0.82,1.10 |
| Richer | 0.71 | 0.59,0.85 | 0.81 | 0.67,0.99 |
| Richest | 0.65 | 0.52,0.82 | 0.55 | 0.43,0.71 |
| Region | | | | |
| North | 0.66 | 0.59,0.75 | 0.75 | 0.66,0.86 |
| Central | Ref | Ref | Ref | Ref |
| East | 0.71 | 0.64,0.79 | 0.61 | 0.54,0.69 |
| Northeast | 0.79 | 0.67,0.94 | 0.68 | 0.57,0.82 |
| West | 0.61 | 0.49,0.77 | 0.47 | 0.36,0.61 |
| South | 0.50 | 0.42,0.60 | 0.49 | 0.40,0.59 |
| Urban residence | | | | |
| Rural | Ref | Ref | Ref | Ref |
| Urban | 0.86 | 0.75,0.98 | 0.95 | 0.82,1.10 |
| **Sibling-related characteristics** | | | | |
| Birth spacing | | | | |
| Less than 2 years since previous birth | Ref | Ref | Ref | Ref |
| 2 years or more since previous birth/1st born | 0.54 | 0.49,0.60 | 0.52 | 0.46,0.58 |
| Older sibling died prior to birth | | | | |
| No | Ref | Ref | Ref | Ref |
| Yes | 1.59 | 1.40,1.80 | 1.58 | 1.40,1.79 |

Our exploratory models stratifying each sibling composition group in NFHS-4 data of all births in the past one to five years offer further evidence of differential associations between sibling composition and infant mortality by sex (Appendix Table 5). Among children with two living older sisters and no brothers, males had significantly lower mortality than females (All children born in the past one to five years: AOR 0.74, 95% CI 0.59−0.93; most recent child: AOR 0.62, 95% CI 0.43−0.88). Among children with a mix of older living siblings, males had significantly lower mortality than females in models limited to the most recent child (AOR 0.82, 95% CI 0.68−0.99).

Our models stratifying by birth order NFHS-4 rather than sibling group again offer evidence of differential associations between sibling composition and infant mortality by sex of infant (Appendix Table 6). Among second order births, males had higher overall odds of mortality than females, but having an older sister conveyed significantly less risk for males than for females (significant among all children born in past one to five years subsample only; AOR on interaction term 0.69, 95% CI 0.56−0.86). Among third-order births, there was no significant difference in odds of mortality by sex or sibling composition alone, but males with two older sisters had lower odds of mortality (All children born in the past one to five years: AOR on interaction term 0.48, 95% CI 0.31−0.74; Most recent child: AOR 0.40, 95% CI 0.21−0.75). There was no differential impact of having one living older brother and one living older sister.

Finally, given the findings from the above described exploratory analyses, particularly with regard to sex differences in risk for infant mortality based on sex of siblings for third-order births, we conducted two additional post-hoc analyses using the most recent wave of data, NFHS-4 (2014−15): 1) wealth-stratified logistic regression analysis with third-order births and 2) summaries of crude likelihoods of dying in infancy among the study population, stratified by sibling composition and sex for third-order births. For the wealth-stratified models, we considered differential risk for infant mortality for boys born after two girls or a boy and a girl, relative to a girl born after two boys (Appendix Table 7). Findings document that sex as a protective factor for boys born after two girls, relative to a girl born after two boys, is significant for only the two highest wealth strata groups, though a clear and consistent gradient of increased protection is seen moving from the poorest to the wealthiest.

In analysis of crude likelihoods of infant deaths by sex in this sample, we found that, although percent of infant deaths in the study population regardless of birth order was significantly higher among males (4.00%, 95% CI 3.79−4.12) than females (3.35%, 95% CI 3.20−3.50), for third-order births, this sex difference was lost (third-order male infant deaths 3.29%, 95% CI 2.93−3.65; third-order female infant deaths 3.50%, 95% CI 3.13−3.87). Further analysis with consideration of sex of older siblings for third-order births found that percent of children dying in infancy did not differ significantly by sex of the index child among third born children with two older brothers (males 4.48%, 95% CI 3.43−5.52; females 3.51%, 95% CI 2.72−4.31) or children with one older brother and one older sister (males 3.74%, 95% CI 3.16−4.32; females 3.35%, 95% CI 2.83−3.87). However, for third born children with two older sisters, females were significantly more likely to die in infancy (3.69%, 95% CI 3.02−4.36) than were males (2.24%, 95% CI 1.82−2.66), a 64% greater risk for death for girls relative to boys in this group.

## 4. Discussion

Findings from this study reinforce prior research documenting that sex differences in infant mortality persist in India in ways that suggest persistence of son preference as a social norm affecting infant survival in India [2−6,12]. Findings extend prior research by documenting that sex differences have not declined significantly over the past 25 years, and are most clearly seen for infants born following two sisters and no brothers. These third-order boys were less likely to die in infancy relative to second born boys with an older brother, as well as relative to girls born following two sisters and no brothers, suggestive of higher family valuation of these boys. This lower

**Table 3**
Multivariate factors associated with infant mortality in most recent birth only, India 2015–2016.

| | Males only (unweighted n = 67,366) | | Females only (unweighted n = 55,332) | |
|---|---|---|---|---|
| | Adjusted odds ratio | 95% confidence interval | Adjusted odds ratio | 95% confidence interval |
| **Sibling Composition** | | | | |
| No living older siblings | 1.46 | 1.13,1.89 | 1.54 | 1.16,2.03 |
| 1 living older brother | *Ref* | *Ref* | *Ref* | *Ref* |
| 1 living older sister | 0.84 | 0.64,1.11 | 0.85 | 0.62,1.18 |
| 2 living older brother, no older sisters | 1.16 | 0.82,1.64 | 1.13 | 0.78,1.65 |
| 2 living older sisters, no older brothers | 0.59 | 0.43,0.82 | 1.08 | 0.77,1.51 |
| Mix of older living brothers & sisters | 0.89 | 0.67,1.18 | 1.29 | 0.95,1.74 |
| **Mother characteristics** | | | | |
| Mother's age at birth | 0.98 | 0.89,1.08 | 0.97 | 0.86,1.10 |
| Squared mother's age at birth | 1.00 | 1.00,1.00 | 1.00 | 1.00,1.00 |
| Mother's age at marriage | 0.99 | 0.97,1.01 | 0.99 | 0.97,1.01 |
| Mother's education | | | | |
| None | *Ref* | *Ref* | *Ref* | *Ref* |
| Any primary | 0.84 | 0.69,1.02 | 0.93 | 0.77,1.13 |
| Some secondary, not completed | 0.74 | 0.57,0.97 | 0.79 | 0.59,1.05 |
| Completed secondary or more | 0.47 | 0.34,0.65 | 0.64 | 0.44,0.93 |
| **Household characteristics** | | | | |
| Household structure | | | | |
| Nuclear | *Ref* | *Ref* | *Ref* | *Ref* |
| Non-nuclear | 1.00 | 0.86,1.17 | 1.04 | 0.88,1.24 |
| Religion | | | | |
| Hindu | *Ref* | *Ref* | *Ref* | *Ref* |
| Muslim | 0.86 | 0.70,1.06 | 0.97 | 0.78,1.20 |
| Christian/ Other / No religion | 1.27 | 0.87,1.84 | 1.08 | 0.73,1.58 |
| Scheduled tribe/scheduled caste | | | | |
| No | *Ref* | *Ref* | *Ref* | *Ref* |
| Yes | 1.03 | 0.88,1.21 | 0.94 | 0.79,1.12 |
| Household wealth quintile | | | | |
| Poorest | *Ref* | *Ref* | *Ref* | *Ref* |
| Poorer | 0.98 | 0.81,1.18 | 0.85 | 0.69,1.05 |
| Middle | 0.83 | 0.66,1.04 | 1.02 | 0.79,1.31 |
| Richer | 0.62 | 0.45,0.85 | 0.68 | 0.47,0.97 |
| Richest | 0.61 | 0.42,0.89 | 0.56 | 0.36,0.89 |
| Region | | | | |
| North | 0.88 | 0.71,1.09 | 0.92 | 0.74,1.13 |
| Central | *Ref* | *Ref* | *Ref* | *Ref* |
| East | 0.90 | 0.75,1.08 | 0.73 | 0.60,0.89 |
| Northeast | 1.26 | 0.97,1.64 | 0.75 | 0.55,1.02 |
| West | 0.85 | 0.60,1.20 | 0.74 | 0.49,1.12 |
| South | 0.84 | 0.63,1.13 | 0.73 | 0.53,1.00 |
| Urban residence | | | | |
| Rural | *Ref* | *Ref* | *Ref* | *Ref* |
| Urban | 0.97 | 0.78,1.20 | 0.78 | 0.60,1.03 |
| **Sibling-related characteristics** | | | | |
| Birth spacing | | | | |
| Less than 2 years since previous birth | *Ref* | *Ref* | *Ref* | *Ref* |
| 2 years or more since previous birth/1st born | 0.71 | 0.60,0.84 | 0.55 | 0.46,0.66 |
| Older sibling died prior to birth | | | | |
| No | *Ref* | *Ref* | *Ref* | *Ref* |
| Yes | 1.76 | 1.47,2.11 | 1.53 | 1.26,1.86 |
| **Healthcare factors** | | | | |
| Received antenatal care visits | | | | |
| None | *Ref* | *Ref* | *Ref* | *Ref* |
| 1–3 | 0.93 | 0.77,1.13 | 1.07 | 0.88,1.30 |
| 4 or more | 0.85 | 0.66,1.08 | 1.00 | 0.79,1.26 |
| Birth had skilled birth attendant | | | | |
| No | *Ref* | *Ref* | *Ref* | *Ref* |
| Yes | 0.93 | 0.76,1.13 | 1.06 | 0.88,1.30 |
| Breastfeeding initiated within 1 h | | | | |
| No | *Ref* | *Ref* | *Ref* | *Ref* |
| Yes | 0.26 | 0.22,0.30 | 0.29 | 0.25,0.34 |

mortality risk may be attributable to greater nutritional and health care investments in these male children, a theory supported by prior research showing that parents in India are more likely to breastfeed and seek required health care for male relative to female children [8–11,18,19]; such behaviors may be even greater for boys born

subsequent to only daughters. There is extensive evidence from India regarding prenatal fertility practices in families with only girls to achieve a more highly desired boy, which likely results in greater investment for that boy and exacerbates the sex ratio imbalance and excess female mortality, with these effects demonstrated in wealthier
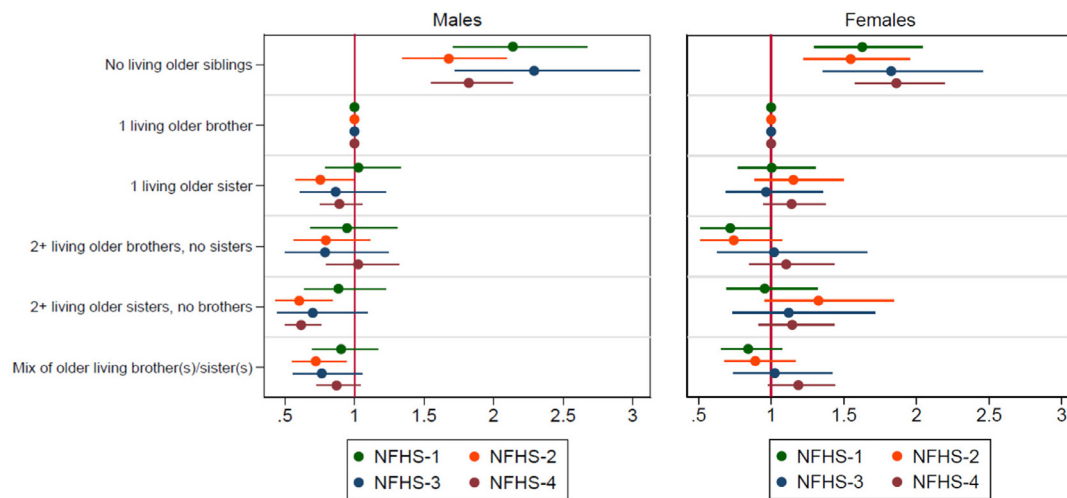
**Fig. 1.** Adjusted odds of infant mortality among all births in prior one to five years by older sibling composition.

more than poorest populations [1,12,20,21]. Our findings also demonstrate stronger protective effects for boys in the wealthier and wealthiest populations. Given that the observed sex differences in sibling effects found in this study remain unchanged for the past two decades, the success of efforts to address son preference in India over this 20-year period must be called into question, and that this resistance to change is seen most heavily in wealthier populations.

Our primary analyses suggest that sex composition of older siblings do not appear to affect female infant mortality, which is in line with other recent research showing that there is no longer increased risk for female relative to male infant mortality in India [5]. However, our exploratory analyses do indicate an increased risk of infant mortality for higher birth order girls (i.e., child three or higher) relative to boys, with third-order girls with two older sisters 64% more likely to die in infancy relative to third-order boys with two older sisters. Correspondingly, the sex ratio IMR for third order infants with two older sisters is 1.4, indicating a 40% higher IMR for males relative to females in this group. These findings are consistent with research from India indicating that as fertility rates decline, differential treatment and survival of higher order infants based on sex worsens [22]. Further, national data also indicate that higher fertility, and resultant higher order births, is more likely among poorer populations in the country [1], and in these populations we continue to see excess female postnatal mortality [3,5]. Overall, findings from this study and prior research suggests that health care outreach to poorer and higher birth order girls must be prioritized to address excess female infant mortality rates, particularly given that risks for postnatal mortality are largely due to treatable infectious diseases [23]. Certainly, India's shift toward smaller family size can help address this issue, though it will not address the underlying norms of son preference and its resultant harms. Notably, in our exploratory analyses focused on observed effects for third order births by wealth quintile found that observed indications of son preference and infant mortality are actually strongest for higher wealth quintiles. Prior research has found sex disparities in post-natal infant mortality are equivalent across socio-economic strata [13]. These findings suggest that focus on supports for third-order girls in India should be a focus regardless of economic strata.

Importantly, these findings as well as other research highlight that neither efforts to address sex ratio imbalance through restricted abortion access nor financial incentives for two girl only families have proven sufficient to tackle sex ratio imbalance in India [3,22], nor have they had any impact on treatment or survival of girls once born [12]. Likely this is because they do not address the underlying social norms that maintain son preference and sex discrimination across India. Interestingly, spatial analyses indicate clustering of postnatal discrimination against girls in India in different regions than those seen to have the most skewed sex

ratios at birth [3]. Hence, addressing norms related to son preference and reduced health care seeking for girls, inclusive of requiring accountability of health systems to outreach to girl infants, may more effectively affect both excess female mortality and sex ratio imbalances than do existing policy efforts. While building a health system response to this issue may be difficult, it is possible in India given the emphasis on household community outreach via ASHAs (community health workers) in India's public health system [24]. ASHAs are incentivized to identify pregnant women at the household level, support women to receive antenatal care and institutional delivery, and so are well-positioned to increase health system awareness and response to households with vulnerable third born girls. To date, consideration of how the health system can respond to differential treatment and survival of infants by sex has not been a priority. This is a major recommendation of the larger Lancet Series on Gender Equality, Norms, and Health to which this paper is attached [25,26]. Importantly, while ASHAs would be instrumental in helping identify vulnerable infants, intervention cannot solely be the responsibility of the already over-burdened ASHAs. The system must respond to known gender inequalities in the same vein as other social inequalities, as part of a commitment to Universal Health Coverage to leave no one behind [25].

Consistent with prior research, first born infants regardless of sex were at increased risk for mortality [6]. Findings from this study did suggest an attenuated effect of being a first-born child on risk for infant mortality once accounting for essential maternal and neonatal care practices. These practices included early breastfeeding initiation, antenatal care (ANC), and having a skilled birth attendant (SBA), all of which have seen dramatic improvement in India from 2005—06 to 2015—16 (breastfeeding initiation within one hour of birth: 23% to 42%; 4+ ANC visits: 37% to 51%; SBA: 47% to 81%) [1]. In adjusted analyses, only breastfeeding was significantly associated with infant mortality. Given that early initiation of breastfeeding was still reported by the minority of participants, promotion of early initiation of breastfeeding requires more attention. Research from India suggests that breastfeeding duration can be shorter for female relative to male children, particularly when there is desire to promote fertility and have a boy child [21]. Hence, sustained support for breastfeeding girls may be needed as well.

Study findings should be considered in light of certain limitations. First, this study relies on self-report and is thus vulnerable to recall bias and social desirability. As the issues of focus (sex of child and siblings, death of child) are limited to past five years and are very basic family demographic data, we assume recall bias and social desirability issues are minimal. Further, NFHS data collection techniques include systems of checks and assessment approaches to help ensure high quality reliable data. An additional limitation is reliance on cross-sectional surveys for a population of women of childbearing

age. The cross-sectional study design limits assumptions of causality in our findings, and the inclusion of women who may still be growing their family restricts assumptions based on completed families. Further threats to a causal interpretation come from the fact that sex-selective abortions are a possibility in India. Hence, the sex of the child after the first birth cannot be considered random. There are also sample selection issues that arise as a result of only looking at sibling composition among alive siblings. While we are unable to effectively deal with these issues given the nature of available data, future research could attempt to address these issues more carefully.

We considered potential confounders based on variables previously seen to be related to this outcome in India [1], but it is likely that we did not adjust for all possible confounding. For example, while low birth weight is a known risk factor for infant mortality [17], this variable was not included in our models due to high rates of missingness across study waves, even in NHFS-4, where 23% of participants did not have birth-weight data. Post hoc analyses suggest that birthweight is associated with infant mortality, and inclusion of birthweight affected the point estimates for our odds ratios for sibling composition variable. However, further examination revealed that estimates changed due to the loss of representativeness of the sample when we dropped the 23% with missing birthweight data. Birthweight missingness was non-random: the IMR for those with birthweight data is 24.5 and the IMR for those without birthweight data is 77.8 (p < 0.001). Given the rapid increase in facility delivery in India, we anticipate that analysis of the next wave of NFHS data will allow for inclusion of birth weight to better consider this confounder. Additionally, while our study uses sex-disaggregated data to understand differential risks for infant mortality by sex as a means of understanding son preference, data on son preference ideologies inclusive of relative value of boy and girl children in a household were not directly assessed. As such, interpretations of the observed findings may be indicative of underlying social norms and attitudes, but we cannot confirm this directly.

There has been extensive research documenting son preference in India for decades, and concern regarding India's sex ratio imbalance is well-recognized and has been cross-validated using data on self-reported desire for more sons than daughters documented at the national level [1]. Nonetheless, there have been shoddy studies which have led to questions regarding the value and integrity of research on sex ratio imbalances [27]. For example, studies suggesting that physiologic traits or toxicity exposure affect likelihood of male births were published but eventually disproven [27]. We feel more confident regarding the validity of our findings because they are consistent with prior research indicating differential risk for infant mortality by sex in India, based on older data [6,12,13]. Hence, these findings highlight continuation of an ongoing and recognized concern. Additionally, our findings related to greater protection for third born boys with older sisters and not brothers, relative to same positioned girls, appear to be very robust, demonstrating significant across models and post-hoc analyses. Multiple testing is a limitation.

## 5. Conclusion

Findings from this analysis of nationally representative data from India indicate that sex composition of older siblings differentially affects risk for infant mortality by sex, with boys born after two sisters and no brothers having lower odds of dying in infancy relative to both boys born after an older brother and girls born after two sisters and no brothers. Further, these sex differences in sibling effects on infant mortality, indicative of son preference, show no decline over the past 25 years, demonstrating that despite recent indications of improvement in equity and infant survival in India [5], son preference persists and continues to contribute to excess female mortality in the country. Time trend analysis of data from 1992—93 to 2015—16 indicate that with reduction in national fertility rates, higher birth order girls regardless of sex composition of older siblings have an increasing risk for infant mortality. These findings correspond with prior studies indicating that decreasing fertility rates are associated with a growth in excess female mortality in India [3,12]. National goals related to the decline in infant mortality will not be achieved unless excess female mortality is addressed [3]. Current policy efforts focused on affecting sex ratio imbalance may have reduced sex ratio imbalance but appear to have had no effect on underlying son preference that sustains this ongoing imbalance [12]. Further efforts will likely require changes in current social norms regarding both son preference and reduced investment in the care of girls, combined with health system outreach and accountability to reach vulnerable girls in India.

## Statement of Authorship

AR led conceptualization and writing on this study, and supported development of the analytic approach. NEJ, LM and AS led development of the analytic approach on this study, and NEJ led data analysis and co-wrote the first draft of the paper with AR. AT, PB, JGS, KK and LL provided substantive input on conceptualization and analytic approach for this paper and helped revise and refine the paper for submission. All authors reviewed this final version of the paper and approved it prior to submission.

## Declaration of Competing Interest

The authors have no conflicts of interest to declare.

## Appendix A.  Supplementary Data

Supplementary data to this article can be found online at https://doi.org/10.1016/j.eclinm.2019.08.016.

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