Journal of Demographic Economics (2021), **87**, 67–106
doi:10.1017/dem.2020.5

CAMBRIDGE UNIVERSITY PRESS

RESEARCH PAPER

Son preference and the fertility squeeze in India

Anna-Maria Aksan

Department of Economics, Fairfield University, 1073 North Benson Road, Fairfield, CT 08624-5195, USA
Corresponding author. E-mail: aaksan@fairfield.edu

(Received 26 August 2018; revised 15 March 2020; accepted 16 March 2020)

Abstract

India's sex ratio at birth (SRB) has become increasingly masculine, coinciding with a declining total fertility rate. Couples often resort to sex-selective abortion or even infanticide to ensure the birth of a son. We use household-level data from all four waves of India's National Family Health Survey (NFHS) to investigate the effect of India's demographic transition on the SRB. Mixed-effects logit regression analysis demonstrates that since the early 1990s, the probability of a third-order birth being male is decreasing in the number of sons previously born, while for second-order births, this effect does not become apparent until the 2000s. Accounting for geographic heterogeneity in the fertility transition, we find additional heterogeneity in son preference and sex selection at the village/neighborhood level. This heterogeneity has strengthened over time at both second- and third-order births and in more and less gender skewed regions of the country, suggesting potential convergence in sex ratios across India. By incorporating the most recent 2015–16 round of the NFHS, we demonstrate that previously documented trends in sex selection continue, and additionally that sex selection is increasingly occurring at lower parities as the desire for a smaller family combines with the traditional preference for sons.

Key words: Fertility decline; fertility squeeze; mixed-effects; sex-selective abortion; son preference

1. Introduction

India has a long history of preference for sons over daughters [Sen (1990), Mayer (1999), Das Gupta (2005), Madan and Breuning (2014)]. A vast literature discusses the economic and cultural factors contributing to this son preference [Arnold *et al.* (1998), Basu (1999), Mayer (1999), Croll (2000), Chung and Das Gupta (2007), Guilmoto (2009), Bhaskar (2011), Jayachandran (2015)]. The spread of ultrasound technology has allowed the practice of sex-selective abortion to proliferate, despite its legal ban in India under the Pre-Natal Diagnostic Techniques Act passed in 1994 [Bhalotra and Cochrane (2010), Madan and Breuning (2014)], magnifying the demographic effects of more traditional methods of gender control such as stopping behavior and even infanticide or neglect of newborn daughters [Das Gupta and Bhat

© Université catholique de Louvain 2021

https://doi.org/10.1017/dem.2020.5 Published online by Cambridge University Press

(1997), Bongaarts (2013)].[1] The most recent India census data of 2011 indicate that the sex ratio at birth (SRB) was 1.12, up from 1.10 in 2001 [Rajan *et al.* (2017)].[2] By some counts, the SRB has reached as high as 1.3–1.5 in parts of Punjab and Haryana, the states with the most skewed SRBs [Guilmoto (2009)]. Although data sources at times conflict, as illustrated in Rajan *et al.* (2017), the most recent estimates provided by the Sample Registration System indicate a continuation of the national trend as recently as 2015 (Census India).[3]

India's growing SRB imbalance coincides with the country's fertility decline. The total fertility rate (TFR) fell from 6 children per woman in the 1950s to 4.7 during 1976–81 to 2.7 in 2005–06 and most recently stands at 2.35 [World Bank Indicators (2015)]. There is considerable variation though, with fertility as high as 4 in some states and below replacement level in others [Dharmalingam *et al.* (2014)]. Fertility is generally higher in India's North where sex ratios are also most distorted, consistent with the view that son preference and large ideal family size are positively correlated [Bongaarts (2013), Jayachandran (2017)]. Although son preference tends to decline with economic development, the sex ratio imbalance tends to worsen initially, as a desire to have a son couples with the desire for a smaller family, the so-called fertility squeeze [Jayachandran (2015)]. The trajectory of India's sex ratio transition has important implications for the country's demographic profile and the well-being of its girls and mothers [Guilmoto (2009), Bongaarts (2013), Kashyap and Villavicencio (2016)].

In this paper, we investigate the dynamics of a fertility squeeze in India using nationally representative, individual-level data from all four rounds of the National Family Health Survey (NFHS) spanning the early 1990s through 2015–16. This is the first paper to combine all four waves of India's NFHS to examine the influence of the fertility transition on the country's notoriously skewed SRB. Specifically, we examine the manifestation of son preference in the form of sex selection at various birth orders, accounting for differences in the timing of the fertility transition throughout India. Regression results here confirm a fertility squeeze effect, with urban areas leading rural ones through the sex ratio transition. Previous empirical studies on India's sex ratio imbalance have been mostly descriptive in nature, or have applied regression analysis only to a small segment of the country. By incorporating the most recent 2015–16 round of the NFHS data, our results confirm previously identified trends and reveal an intensifying fertility squeeze, contributing to the literature examining the sex ratio imbalance. Specifically, sex selection continues to occur at third-order births but more recently is also evident at second-order births. A re-balancing of India's sex ratio is not expected for the near future.

Substantial variation in SRBs throughout India [Bongaarts (2013)] demonstrates differences in son preference as well as ability and willingness of parents to employ sex-selective abortion in order to achieve their ideal family composition. While

[1]Abortions are legal but restricted in India, as dictated by the Medical Termination Act of 1971 [Rajan *et al.* (2017)]. Amniocentesis was first introduced to India in the mid-1970s, leading to immediate increases in abortion of female fetuses, and ultrasound technology was introduced a few years later, both spreading throughout the 1980s [Guilmoto (2009)].

[2]The overall sex ratio did improve 2001–2011; however, from 1.07 to 1.06 due to improved longevity among women [Dandona and India State-Level Disease Burden Initiative Collaborators (2017), Madan and Breuning (2014)].

[3]According to the Sample Registration System, the SRB for the country has worsened from 1.10 in 2012–2014 to 1.11 in 2013–2015.

https://doi.org/10.1017/dem.2020.5 Published online by Cambridge University Press

access to ultrasound technology is widespread in India, it remains illegal to learn the fetal sex, and this barrier may be easier to overcome in some areas. Regardless, son preference will lead some parents to engage in stopping behavior while others will use sex-selective abortion to manipulate sex ratios. Past studies have investigated the correlation of sex ratio distortions with religion, age, education levels, socioeconomic status, and family composition. The analysis here accounts for these factors, but also allows for additional heterogeneity in the tendency to sex select by estimating mixed-effects models, which allow for different responses to the sex composition of one's existing children. Mixed-effects also account for some of the estimation bias inherent in modeling sex selection by controlling for unobserved heterogeneity.[4] This technique has had only limited applications in studying sex ratio imbalances [Bhalotra and Cochrane (2010)], and this paper is the first to investigate heterogeneity at the community level, reflecting the influence of social norms on both desired fertility and sex ratio ideals.[5] Unlike Bhalotra and Cochrane (2010) who conclude that heterogeneity at the individual mother level is of little importance in India, we find growing unobserved heterogeneity in sex selection at the community level, in the notoriously gender skewed Northwest but even more so in the rest of the country where sex ratios have historically been more balanced [Das Gupta and Bhat (1997)]. This suggests convergence across India, although toward a masculine rather than balanced SRB.

2. Background and literature

As parents choose to have fewer children, the probability of ending up sonless increases, exerting pressure to sex select when there is a strong preference for sons. This results in a "fertility squeeze" on the SRB [Guilmoto (2009)]. Traditional methods, in particular fertility stopping behavior whereby parents have children until the desired number of sons is born, distort SRBs at the individual family level, with girls generally growing up in larger families, but not at the national level [Bongaarts (2013)]. The "stopping rule" assumes that sonless families are willing to keep having children, but as desired fertility falls, this approach becomes less effective in achieving multiple family planning goals. With the widespread availability of non-invasive, prenatal sex-determination technology in the form of the ultrasound, parents can effectively choose to have more sons and fewer daughters. For example, falling fertility combined with son preference led to rapid increases in SRBs during the 1990s in all newly independent countries of the South Caucasus, reaching 1.17 in Azerbaijan in 2002, 1.19 in Georgia in 1998, and 1.16 in Armenia in 2001 [Guilmoto (2009)]. Similarly, China's TFR fell from 2.9 in 1979, just prior to the implementation of the one-child policy, to 1.7 in 2004, while the sex ratio worsened from 1.06 in 1979 to 1.17 in 2001 [Hesketh *et al.* (2011)]. Yang (2012) illustrates a clear fertility squeeze in China, with populations subject to the 1-child policy resulting in the most severely imbalanced SRB at first-order births, those subject to the 1.5-child policy at second-order births, and those subject to a 2-child limit regardless of the sex of first born exhibiting the

[4]The various sources of bias are discussed in Section 4.1.

[5]See, for instance, Munshi and Myaux (2006) who demonstrate distinct fertility behaviors across religious groups within villages in rural Bangladesh.

https://doi.org/10.1017/dem.2020.5 Published online by Cambridge University Press

most balanced sex ratios [Li *et al.* (2011)].[6] India's fertility decline has been slower and so sex selection may not manifest until higher order births, but a similar fertility squeeze is expected [Das Gupta and Bhat (1997), Bhalotra and Cochrane (2010), Dharmalingam et al. (2014), Kashyap and Villavicencio (2017)].

India's accelerated fertility decline began in the 1980s and initially appeared to exacerbate regional differences in sex ratio imbalances. The rise in the imbalance has been greater in the North than in the South and has always been highest in the Northwest; correspondingly fertility declined more rapidly in the Northwest than the rest of the North [Das Gupta and Bhat (1997)]. On the other hand, fertility has declined most rapidly in the South but the sex ratio has risen least there, so historical regional differences in sex discrimination persist [Das Gupta and Bhat (1997)]. Yet Basu (1999) describes the surprising skewing of SRBs in traditionally more gender equal southern India as already low fertility has fallen even further, while some northern states have demonstrated a reversal in SRB imbalances alongside falling fertility.[7] More recent analysis by Diamond-Smith and Bishai (2015) demonstrates that states with the most skewed SRBs have begun to reverse the course. These reversals suggest possible convergence rather than a clear weakening or strengthening of son preference as India's demographic transition proceeds [Rajan *et al.* (2017)].

However, apparent convergence may be partially attributed to a mathematical phenomenon rather than the implied changes in behavior [Kashyap and Villavicencio (2016), Dubuc and Sivia (2017)]. At low fertility, a small proportion of sex-selective procedures suffices to significantly distort the SRB. Importantly, this "disproportionality effect" illustrates how SRBs can be an inaccurate measure of son preference, and can explain some of the inconsistencies demonstrated in the empirical literature between falling son preference but rising sex ratios during the fertility transition [Bongaarts (2013)]. Rather than focusing on aggregate sex ratio measures, this paper examines the extent to which sex of a birth can be predicted by the sex composition of previous births, in line with several previous studies.8 Our regression results demonstrate continuing sex selection in the Northwest but with the addition of selection at increasingly earlier parities, as well as some emerging evidence in the rest of the country most recently. Thus at the very least, further masculinization of the SRB is expected via the disproportionality effect.

Many descriptive empirical studies illustrate India's fertility squeeze as shrinking family size further skews SRBs [Croll (2002), Guilmoto (2009, 2012)]. Like our

[6]The one-child policy allowed for many exceptions, among others, some families were permitted to have a second child if their first child was a girl.

[7]Basu identifies three regimes: (1) high fertility/high son preference in the North recently and some states today, (2) medium fertility/medium son preference in the South recently and some northern states today, and (3) low fertility/continued medium son preference in some states such as Tamil Nadu in the South. Kerala represents a possible fourth regime of low fertility/low son preference.

[8]Arnold *et al.* (2002) demonstrate a fertility squeeze at the third birth in India based on 1998–99 data. Das Gupta (2005) discusses SRBs at higher parities based on sex composition of existing children. Jha *et al.* (2011) document rising SRBs at second-order births when the first-born is female between 1991 and 2011. Patel *et al.* (2013) demonstrate a higher probability of a son being born if previous children were female among a group of Indian physicians, more so than for the general population. In analyzing China's one-child policy, Qiao (2008) finds that when the first child was a son, the sex ratio of the second birth was 1.01, while it was 1.26 if the first child was a daughter [Yang (2012)].

https://doi.org/10.1017/dem.2020.5 Published online by Cambridge University Press

study, Jayachandran (2017) is one of the few to apply regression analysis. Using cross-sectional survey data from several districts within the state of Haryana, notorious for its unusually skewed sex ratio, Jayachandran demonstrates a sharply rising desired sex ratio as desired fertility falls and estimates that fertility decline may explain one-third to one-half of India's recent sex ratio increase. She examines changes in the desired sex ratio as the hypothetical number of children parents' children might have is varied exogenously. Yet Bongaarts (2013) shows significant differences between desired and actual SRBs and that desired SRBs become more balanced even while actual SRBs become more masculine at lower levels of fertility, based on data from India and other countries. Bhat and Zavier (2003) demonstrate that while the percentage of couples preferring more sons than daughters drops with ideal family size in India, improved access to sex-selective technologies is fulfilling unmet demand thereby raising SRBs in northern India. Jayachandran (2017) also demonstrates that son preference and ideal family size are positively correlated in India, but that the desired sex ratio *rises* sharply as fertility falls. Thus, there is some inconsistency in the literature regarding the relationship between desired sex ratios and fertility change.

We instead consider geographic variation in fertility and examine its impact on the sex composition of *actual* births. We use a larger nationally representative data set that allows for regional comparisons across India, while Jayachandran (2017) focuses on one state with a particularly masculine SRB. Jayachandran uses OLS to estimate the relationship of percent of sons desired and desired family size and uses the results to calculate the overall contribution of fertility decline to rising sex ratios in India. We estimate the probability of a male birth conditional on (a different measure of) desired fertility as well as the sex composition of previous births, interacting the two to capture the fertility squeeze. We demonstrate a tightening of the squeeze over time and, by using mixed-effects modeling, growing heterogeneity in sex selection over time.

To measure fertility norms, we use the community-level average number of children born to women at all stages of their fertility. Desired family size is affected by own experiences as well as the behavior of one's age cohort. For instance, individuals who were raised in a large family may be inclined toward having more children [Jennings *et al.* (2012), Kolk *et al.* (2014)]. In addition to the intergenerational transfer of fertility, social network effects also influence fertility decisions [Madhavan *et al.* (2003), Munshi and Myaux (2006), Balbo and Barban (2014), Bernardi and Klärner (2014)]. Couples may choose to have a third child if other couples around them are doing so, but not if most stop at two children. Our measure of local fertility norms incorporates both factors. It also resembles the definition of fertility squeeze in Kashyap and Villavicencio's (2016) simulation model, where the aggregate TFR is used to measure the intensity of the fertility squeeze, and parameter values determine the relationship between this squeeze and actual sex selection. Our specification differs, though, in that the fertility squeeze varies by location and we allow directly for variation in actual sex selection via mixed-effects modeling rather than tying the variation to the fertility squeeze per se.

The paper proceeds as follows. Section 3 introduces the data used in our analysis and presents some descriptive statistics indicating sex selection in India as well as increased variation in this practice over time. Section 4 describes the econometric technique and discusses the regression results. Section 5 concludes with some policy implications.

https://doi.org/10.1017/dem.2020.5 Published online by Cambridge University Press

3. Data

The NFHS is a large-scale, multi-round survey conducted among a nationally representative sample of households covering over 99% of India's population. The survey collected information on population, health, and nutrition, with a focus on women and young children, and included questions about reproductive health and family planning services (World Bank). There have been four rounds of the NFHS for the years 1992–1993, 1998–1999, 2005–2006, and 2015–2016. NFHS-1 sampled 89,777 ever-married women aged 13–49, NFHS-2 sampled 90,303 ever-married women aged 15–49, NFHS-3 sampled 124,385 ever-married women aged 15–49, and NFHS-4 sampled 699,686 women aged 15–49.[9]

In each state, the rural sample was selected in two stages, with the selection of Primary Sampling Units (PSUs), which are villages, with probability proportional to population size (PPS) at the first stage, followed by the random selection of households within each PSU in the second stage. Villages with fewer than five households were omitted, and villages with fewer than 50 households were combined with other contiguous villages. In urban areas, a three-stage procedure was followed. In the first stage, wards were selected with PPS sampling. In the next stage, one census enumeration block (CEB) was randomly selected from each sample ward. In the final stage, households were randomly selected within each selected CEB. An average of 20–30 households were targeted per PSU with a target range of 15–60.

Reflecting changing geopolitical boundaries in India over time, the first survey covered 25 states, the second and third 26 states, and the fourth 29 states as well as India's six union territories. Newly created states may differ from their mother state in ways that would affect regression results, so we use the newest state and territory definitions when controlling for state fixed effects. However, results are confirmed when aggregating observations to their original state groupings (and dropping observations from the union territories which do not appear in the first three rounds of the NFHS), and are available upon request.

This study uses all four rounds of the NFHS and includes data on all women who have had at least two or three births, depending on the regression model. All descriptive statistics presented below are based on sampling-weighted estimates, since oversampling in some areas could otherwise bias outcomes. The sampling weights are the national-level weights provided by NFHS, which are appropriate when conducting analysis above the state level.

Figures 1 and 2 present the distributions of sampling-weighted PSU-level averages of key variables for each of the four survey years to demonstrate changes in mean and variance over time. All available observations are included in the graphs, not just those used in the final regression samples. In Figure 1, the average number of children born to women in the sample declines over time, reflecting falling fertility since the early 1990s, with a clear leftward shift of the distribution between the 1990s and 2000s. Figure 2 shows that among respondents with a first-born daughter, the second birth is, on average, more likely to be male. The distribution has migrated

[9]While the fourth wave of the NFHS sampled many more women than previous waves, the dataset is missing a substantial number of observations with regards to an important control variable used in the regression analysis below, namely education of the respondent's partner. Thus the final sample used here will be similar in size across all four waves. Regression results omitting partner's education, based on a much larger sample for wave four, yield qualitatively similar results and, for some specifications, stronger statistical significance. These results are available upon request.

https://doi.org/10.1017/dem.2020.5 Published online by Cambridge University Press

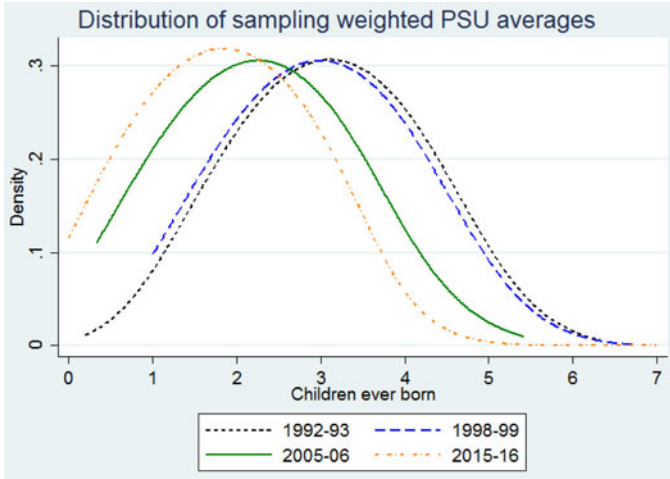


Figure 1. Children ever born per respondent, by survey wave, NFHS.
Note: First PSU averages are calculated, then probability sampling weights are applied when generating the density distributions. All available data are used for women of all ages 15–49.

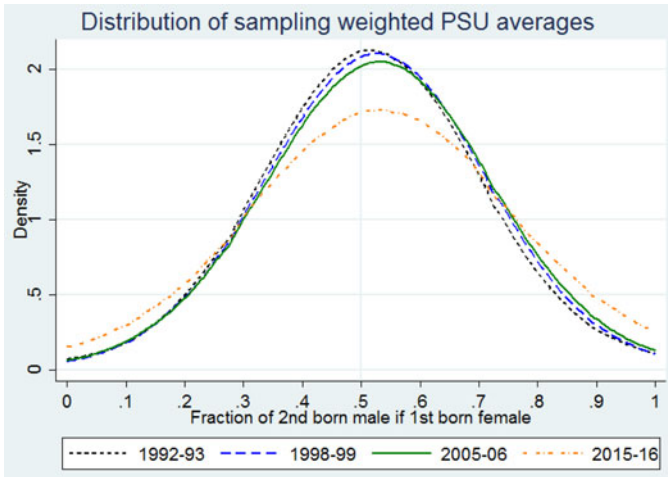


Figure 2. Proportion of second births male if first birth female, by survey wave, NFHS.
Note: First PSU averages are calculated, then probability sampling weights are applied when generating the density distributions. All available data are used for women of all ages 15–49.

further to the right over time, coinciding with the fertility change in Figure 1 and thereby suggesting a fertility squeeze. Moreover, the distribution has widened over time, especially for the most recent sample year, indicating growing variation across PSUs in the practice of sex selection at second-order births. Similar trends can be shown for male third-order births when the first two births are female. Declining fertility accompanied by increasing proportions of male births following female

https://doi.org/10.1017/dem.2020.5 Published online by Cambridge University Press

Table 1. Prevalence of male births by birth order, by regression sample, and by survey wave, NFHS

	1992–93	1998–99	2005–06	2015–16
Sample with 2+ births				
1st born male	0.519	0.518	0.509	0.514
2nd born male	0.519	0.522	0.521	0.511
Sample with 3+ births				
1st born male	0.508	0.499	0.484	0.467
2nd born male	0.509	0.505	0.493	0.462
3rd born male	0.516	0.522	0.526	0.531

Note: Observations are weighted using sampling probability weights.

births suggests a fertility squeeze effect. Parents are feeling growing pressure to sex select as they approach their desired family size.

Table 1 demonstrates the average prevalence of male births by birth order for each sample year. Data are restricted to women included in the final regression samples. Among women with at least two births, more boys are born than girls at every birth order in each sample year, but there is no clear upward or downward trend over time or across births.10 Among women with at least three births, the prevalence of male births is increasing by birth order in every sample year. In 1992, more boys than girls were born, on average, at first, second, and third births, but in 1998, there is a higher prevalence of daughters at first order, and in 2005 and 2015, a higher prevalence of daughters at first and second orders. Increasingly over time, respondents who have had at least three births tend to be those who have more daughters at earlier birth orders, with a clear reversal toward male births at the third order. This suggests an intensifying fertility squeeze whereby higher order births are pursued when the respondent has yet to yield a son.

Tables 2 and 3 present the summary statistics for the samples used in the regression analysis that follows, broken down according to the sex of the respondent's second- and third-born child, respectively. The manifestation of sex selection at second births would present as a lower proportion of first-born male children among respondents who had a son at second birth compared to those who had a daughter at second birth. In 1992, the opposite is true, although the difference is statistically insignificant.[11] For all following years, second-born females are more likely than second-born males to have an older brother, with the difference in means becoming statistically significant for 2005 and 2015. Regardless of year and sex of the second child, it is more common for the first child to be male, i.e., both female and male second births are more likely to have an older brother than an older sister. Fertility has clearly declined over time and is on

[10]While a slightly higher proportion of males tends to be born even in populations not practicing sex selection, approximately 105 male per 100 female births, the sex ratio eventually balances out due to naturally higher male infant and child mortality. Notably, neglect of daughters remains a widely used method of sex selection so that under-5 mortality is actually higher among girls than boys in Nepal, Pakistan, India, and especially China [Guilmoto (2009)].

[11]Specifically, among second-born females, 51.7% have an older brother and thus 48.3% have an older sister, while for second-born males, 52.0% have an older brother and thus 48.0% have an older sister.

https://doi.org/10.1017/dem.2020.5 Published online by Cambridge University Press

https://doi.org/10.1017/dem.2020.5 Published online by Cambridge University Press

Table 2. Summary statistics of a sample of second-born children by sex of the child and by survey wave, NFHS

	1992		1998		2005		2015	
	Female	Male	Female	Male	Female	Male	Female	Male
1st born male	0.517	0.520	0.520	0.516	0.518**	0.501**	0.523**	0.505**
	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.499)	(0.500)
Children ever born	4.119**	3.925**	3.924**	3.701**	3.763**	3.479**	3.213**	2.913**
	(1.960)	(1.904)	(1.862)	(1.815)	(1.789)	(1.722)	(1.443)	(1.295)
1st child died	0.166	0.169	0.144	0.149	0.136	0.134	0.0810	0.0808
	(0.372)	(0.375)	(0.351)	(0.356)	(0.343)	(0.341)	(0.273)	(0.273)
Mother's education	2.585	2.553	3.070	3.043	3.538	3.590	5.033**	5.135**
	(4.028)	(4.008)	(4.237)	(4.244)	(4.484)	(4.565)	(4.918)	(4.942)
Partner's education	5.371	5.304	5.847	5.747	6.031	6.084	6.860**	7.000**
	(4.995)	(5.009)	(4.976)	(5.018)	(5.076)	(5.109)	(4.920)	(4.967)
Household head	0.0511*	0.0471*	0.0527*	0.0486*	0.0964**	0.0858**	0.102*	0.0962*
	(0.220)	(0.212)	(0.223)	(0.215)	(0.295)	(0.280)	(0.302)	(0.295)
Wife of household head	0.683	0.683	0.689	0.692	0.718	0.720	0.684**	0.673**
	(0.465)	(0.465)	(0.463)	(0.462)	(0.450)	(0.449)	(0.465)	(0.469)
Daughter of household head	0.0465	0.0457	0.0461	0.0446	0.0164	0.0155	0.0218	0.0214
	(0.211)	(0.209)	(0.210)	(0.206)	(0.127)	(0.124)	(0.146)	(0.145)
Daughter-in-law of household head	0.149	0.148	0.151	0.150	0.138	0.145	0.167**	0.179**
	(0.356)	(0.355)	(0.358)	(0.357)	(0.344)	(0.352)	(0.373)	(0.384)
Hindu	0.817	0.820	0.818	0.815	0.809	0.811	0.805**	0.814**
	(0.387)	(0.384)	(0.386)	(0.388)	(0.393)	(0.391)	(0.396)	(0.389)
Muslim	0.123	0.120	0.127	0.128	0.135	0.134	0.140*	0.134*
	(0.328)	(0.326)	(0.333)	(0.334)	(0.342)	(0.341)	(0.347)	(0.341)

(*Continued*)

https://doi.org/10.1017/dem.2020.5 Published online by Cambridge University Press

Table 2. (*Continued.*)

	1992		1998		2005		2015	
	Female	Male	Female	Male	Female	Male	Female	Male
Christian	0.0248	0.0227	0.0238	0.0250	0.0232	0.0215	0.0269**	0.0231**
	(0.156)	(0.149)	(0.153)	(0.156)	(0.151)	(0.145)	(0.162)	(0.150)
Scheduled caste	–	–	0.183	0.187	0.191	0.196	0.201	0.197
			(0.387)	(0.390)	(0.393)	(0.397)	(0.400)	(0.398)
Scheduled tribe	–	–	0.0874	0.0874	0.0843	0.0840	0.0956	0.0919
			(0.282)	(0.282)	(0.278)	(0.277)	(0.294)	(0.289)
Other backwards class	–	–	0.328	0.329	0.401	0.398	0.448	0.450
			(0.469)	(0.470)	(0.490)	(0.490)	(0.497)	(0.497)
Mother's age	33.15	33.25	33.35	33.53	34.25	34.27	35.79†	35.69†
	(8.019)	(8.020)	(7.928)	(7.965)	(7.732)	(7.701)	(7.570)	(7.586)
Urban	0.261	0.260	0.254	0.259	0.300	0.300	0.334	0.331
	(0.439)	(0.439)	(0.436)	(0.438)	(0.458)	(0.458)	(0.471)	(0.471)
Electricity	0.523	0.522	0.603	0.603	0.668*	0.681*	0.889	0.892
	(0.499)	(0.500)	(0.489)	(0.489)	(0.471)	(0.466)	(0.314)	(0.311)
Radio	0.415	0.420	0.392	0.397	0.312	0.317	0.0840	0.0798
	(0.493)	(0.494)	(0.488)	(0.489)	(0.463)	(0.465)	(0.277)	(0.271)
TV	0.219	0.223	0.353	0.354	0.456*	0.467*	0.679**	0.690**
	(0.413)	(0.416)	(0.478)	(0.478)	(0.498)	(0.499)	(0.467)	(0.463)
No toilet	0.702	0.704	0.652	0.653	0.571	0.574	0.399	0.395
	(0.457)	(0.456)	(0.476)	(0.476)	(0.495)	(0.494)	(0.490)	(0.489)
Observations	31,247	33,530	31,812	34,783	31,409	34,405	30,954	32,802

Note: Observations are weighted using sampling probability weights. Statistically different means are indicated at the †10%, *5%, and **1% levels.

https://doi.org/10.1017/dem.2020.5 Published online by Cambridge University Press

Table 3. Summary statistics of a sample of third-born children by sex of the child and by survey wave, NFHS

	1992		1998		2005		2015	
	Female	Male	Female	Male	Female	Male	Female	Male
1st 2 born male	0.261	0.263	0.259	0.254	0.245	0.239	0.236**	0.213**
	(0.439)	(0.440)	(0.438)	(0.436)	(0.430)	(0.427)	(0.424)	(0.409)
1st 2 born—1 male and 1 female	0.492	0.495	0.492	0.490	0.495	0.493	0.490**	0.476**
	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.500)	(0.499)
Children ever born	4.783**	4.596**	4.643**	4.403**	4.555**	4.256**	4.140**	3.799**
	(1.790)	(1.774)	(1.732)	(1.687)	(1.678)	(1.613)	(1.379)	(1.248)
1st 1 or 2 children died	0.279	0.280	0.2642	0.263	0.241	0.235	0.190*	0.184*
	(0.449)	(0.449)	(0.441)	(0.440)	(0.428)	(0.424)	(0.392)	(0.388)
Mother's education	1.985	1.975	2.197**	2.292**	2.341**	2.507**	3.236**	3.391**
	(3.434)	(3.458)	(3.551)	(3.611)	(3.665)	(3.804)	(4.093)	(4.241)
Partner's education	4.857	4.872	5.153*	5.243*	5.185	5.285	5.664**	5.842**
	(4.810)	(4.774)	(4.788)	(4.793)	(4.816)	(4.856)	(4.705)	(4.788)
Household head	0.0543*	0.0503*	0.0522	0.0554	0.103	0.0998	0.117	0.121
	(0.227)	(0.219)	(0.223)	(0.229)	(0.304)	(0.300)	(0.321)	(0.326)
Wife of household head	0.729	0.728	0.740*	0.733*	0.756	0.752	0.719	0.713
	(0.445)	(0.445)	(0.439)	(0.442)	(0.429)	(0.432)	(0.449)	(0.452)
Daughter of household head	0.0320*	0.0290*	0.0309	0.0310	0.0121	0.0100	0.0148*	0.0128*
	(0.176)	(0.168)	(0.173)	(0.173)	(0.110)	(0.0995)	(0.121)	(0.112)
Daughter-in-law of household head	0.119	0.122	0.119	0.118	0.0986**	0.108**	0.124	0.125
	(0.323)	(0.327)	(0.324)	(0.323)	(0.298)	(0.310)	(0.330)	(0.330)
Hindu	0.815	0.814	0.811	0.811	0.798	0.803	0.778**	0.797**
	(0.389)	(0.389)	(0.392)	(0.392)	(0.401)	(0.398)	(0.416)	(0.402)
Muslim	0.130	0.129	0.139	0.138	0.154	0.149	0.177**	0.162**

(*Continued*)

https://doi.org/10.1017/dem.2020.5 Published online by Cambridge University Press

Table 3. (*Continued.*)

	1992		1998		2005		2015	
	Female	Male	Female	Male	Female	Male	Female	Male
	(0.337)	(0.335)	(0.346)	(0.345)	(0.361)	(0.356)	(0.381)	(0.369)
Christian	0.0210	0.0213	0.0212	0.0200	0.0174	0.0177	0.0208**	0.0155**
	(0.143)	(0.144)	(0.144)	(0.140)	(0.131)	(0.132)	(0.143)	(0.123)
Scheduled caste	–	–	0.198	0.198	0.210	0.205	0.223	0.226
			(0.399)	(0.398)	(0.407)	(0.404)	(0.416)	(0.418)
Scheduled tribe	–	–	0.0963**	0.0898**	0.0983*	0.0917*	0.110	0.107
			(0.295)	(0.286)	(0.298)	(0.289)	(0.312)	(0.309)
Other backwards class	–	–	0.328	0.331	0.406	0.412	0.440	0.446
			(0.470)	(0.470)	(0.491)	(0.492)	(0.496)	(0.497)
Mother's age	34.80	34.89	35.00*	35.13*	35.74	35.80	37.57	37.56
	(7.574)	(7.533)	(7.548)	(7.474)	(7.236)	(7.319)	(7.073)	(7.062)
Urban	0.240	0.238	0.226**	0.235**	0.262	0.261	0.266**	0.284**
	(0.427)	(0.426)	(0.418)	(0.424)	(0.440)	(0.439)	(0.442)	(0.451)
Electricity†	0.497	0.500	0.567**	0.577**	0.616**	0.630**	0.850	0.848
	(0.500)	(0.500)	(0.495)	(0.494)	(0.486)	(0.483)	(0.357)	(0.359)
Radio	0.386	0.393	0.366	0.364	0.285	0.291	0.0743	0.0755
	(0.487)	(0.488)	(0.482)	(0.481)	(0.451)	(0.454)	(0.262)	(0.264)
TV	0.194	0.196	0.311**	0.322**	0.396**	0.414**	0.585*	0.593*
	(0.396)	(0.397)	(0.463)	(0.467)	(0.489)	(0.492)	(0.493)	(0.491)
No toilet	0.732	0.736	0.697	0.690	0.637	0.635	0.484	0.481
	(0.443)	(0.441)	(0.459)	(0.462)	(0.481)	(0.481)	(0.500)	(0.500)
Observations	23,247	24,877	22,705	24,832	20,259	22,486	17,681	19,624

Note: Observations are weighted using sampling probability weights. Statistically different means are indicated at the †10%, *5%, and **1% levels.

average lower among respondents having a second-born son, suggesting that either women who successfully have a son are less likely to continue having children, or conversely that where desired family size is smaller, the pressure to produce a son is stronger. Among women who have had at least three births (Table 3), respondents with a third-order male birth have a lower prevalence of previous male births, with the exception of the 1992 sample when a fertility squeeze effect would have been weakest at this earlier stage of the demographic transition.[12] The difference in means is not, however, statistically significant until 2015.

Previous research suggests more masculine SRBs among the more educated and wealthy in India [Madan and Breuning (2014), Portner (2016)]. More educated women may have a higher opportunity cost of children and thus lower desired fertility, which would affect the birth order at which the pressure to sex select occurs [Clark (2000)]. More educated women may also be better informed about accessing sex-selective abortions [Arokiasamy (2007), Bhat and Zavier (2003)]. On the other hand, greater education of women might be associated with the higher perceived status of women, which would weaken the preference for sons, although this does not seem to be the case in India. In Tables 2 and 3, education has increased over time, with some narrowing of the gap between men and women, but is not consistently higher or lower among families with second- or third-born male children. More homes have electricity, toilets, and a television, indicating improving living standards, while radio ownership has decreased over time. Of these, only television ownership is consistently higher in homes with a second- or third-born male, although the difference in means is statistically significant only more recently. The respondent's relationship to the head of the household includes the categories: self, wife, daughter, daughter-in-law, granddaughter, mother, mother-in-law, sister, other relative, adopted/foster child, not related, sister-in-law, niece, and domestic servant. Only statistics for the first four are presented in Tables 2 and 3. Women residing with their in-laws may face more pressure to sex select [Robitaille (2013)], although this is not apparent in the descriptive statistics. In Table 2, second-born females are more likely to have a mother that is the household head. Although Hinduism has been more strongly associated with son preference than Islam and Christianity, Tables 2 and 3 do not reflect this until 2015 when the differences in means become statistically significant [Guilmoto (2009)]. Religious composition remains stable, with approximately 81% of the Indian population Hindu, 13% Muslim, and 2% Christian in Table 2. In Table 3, a decreasing proportion of families with at least three births are Hindu and an increasing proportion are Muslim, reflecting differences in the pace of fertility decline. The descriptive statistics thus far suggest that the sex of later births is largely conditional on the sex of previous births and is related to fertility levels [Gu and Roy (1995), Yang and Wang (2006)].

4. Empirical analysis

4.1 Methodology

The demand for sex selection may vary at different birth orders depending on desired fertility size such that a family desiring a total of two children would act on their son

[12]In 1992, among third-born girls 26.1% + 49.2% = 75.3% had at least one older brother, while among third-born boys 26.3% + 49.5% = 75.8% did. By 2015, 72.6% of third-born girls and only 68.9% of third-born boys had at least one older brother.

https://doi.org/10.1017/dem.2020.5 Published online by Cambridge University Press

preference at the second birth, while a family desiring a total of three children might not do so until the third birth, in the case that the previous births were female. The following regression analysis accounts for differences in desired fertility by investigating whether the impact of sex composition of children on the sex of a subsequent birth is affected by local fertility rates. We expect the evidence of sex selection to be stronger among women living in areas with lower fertility rates [Guilmoto (2009)]. The use of local fertility rates avoids reporting bias inherent in individually-reported ideal family size [Jayachandran (2017)] and reflects the influence of social norms and geographically-shared socioeconomic influences on fertility [Bloom *et al.* (2008)].

We further explore unobserved geographic heterogeneity in the degree of son preference and sex selection through the use of mixed-effects models. Assuming that retrospective reporting may bias stated ideal sex ratios of children and also that survey respondents do not accurately report if they have engaged in sex selection, the data set lacks accurate information about son preference and sex selection. Mixed-effects models allow for standard random effects (different intercepts for each geographic area) but also for different effects of selected covariates. In our case, the probability of having a son based on the sex composition of previous births is allowed to differ by village or urban neighborhood. This methodology accounts for some of the potential bias inherent in the regression models.

We implement a mixed-effects logit estimation with two levels. At the first level, the indicator $\mathcal{S}on_{ij}$ takes the value of one if woman i in geographic area j gave birth to a son. Otherwise $\mathcal{S}on_{ij}$ takes the value of zero. Second- and third-order births are separately considered: the first set of models considers women who have had at least two births, and the second set of models includes only those women who have had at least three births, estimating the probability of a male second- or third-order birth, respectively. In the second level, we account for geographic heterogeneity by allowing some coefficients to vary across PSU. The two-level logit models are specified as follows:

$$logit(p_{ij}) = X_{ij}\beta + Z_{ij}u_j + \varepsilon_{ij}, \tag{1}$$

where $p_{ij} = \mathcal{P}r(\mathcal{S}on_{ij} = 1)$ for woman i living in PSUj, X is a vector of variables potentially associated with son preference, Z is a subset of factors assumed to have heterogeneous effects across PSUs, and ε is the error term that follows a logistic distribution. β is the vector of relevant fixed effect coefficients to be estimated, and u_j is a set of random effects that depict heterogeneity across PSUs. The random-intercept logit model is obtained when vector Z is reduced to a vector of ones. In this model, the effects of covariates on the probability of a son being born are assumed to be fixed, but the intercept is allowed to vary across PSUs to account for heterogeneity. As an extension of this model, a subset of variables is included in vector Z to estimate a random-coefficient logit model in which the estimated effects of those variables are allowed to randomly vary across PSUs. We explore the possibility that the sex composition of previously born children can have mixed (fixed and random) effects on the probability of a son being born by including sex composition as an indicator variable in vector Z. Finally, robust standard errors are clustered by PSU.

The vector X includes variables that indirectly capture household wealth, education, and exposure to media, while maximizing sample size. Specifically, X includes the presence of electricity and a toilet in the home, education of the woman and that of

https://doi.org/10.1017/dem.2020.5 Published online by Cambridge University Press

her partner, and ownership of a radio or television. We control for woman's age since the reported second- or third-order births could have occurred as recently as in the past year or up to 34 years ago for any particular survey year. Death of a child is included as this could influence fertility decisions within the household. Indicator variables for the most dominant religions are also included in X, since these will reflect differences in traditions which may influence son preference. The respondent's relationship to the household head controls for family pressures in the decision to sex select. An indicator variable for urban is included, since desired family size tends to be lower in urban areas thereby affecting the rate of fertility squeeze. State fixed effects control for state-level heterogeneity in son preference not captured by the individual control variables.

While the caste system permeates Indian society, this is not explicitly included in the main regressions due to inconsistency in reporting across survey years. Many mechanisms by which caste might influence both fertility and son preference, however, are accounted for through the other control variables. Moreover, recent research by Borker *et al.* (2018) demonstrates substantial variation in sex selection within castes in India, so that members of any particular caste are not homogeneously more or less likely to sex select. Nevertheless, we conduct robustness checks of all regressions by including caste for the three years available, 1998–99, 2005–06, and 2015–16.

The rate of fertility decline has differed throughout India [Das Gupta and Bhat (1997)]. The PSU average level of children born per respondent is included in $\mathcal{X}$ as a proxy for the local fertility rate as this may affect fertility decisions at the household level. If fertility rates have declined locally, social norms may push couples to choose to have fewer children and, through the fertility squeeze, to sex select at lower order births. The regressions that follow present results with and without this PSU average number of children born per woman, as well as with the average interacted with the sex of previously born children. To make the PSU average a meaningful measure, and to avoid endogeneity at the individual respondent level, all observations in PSUs numbering less than five households are dropped from the regression analysis.[13] These make up an imperceptible fraction of the observations, 71 out of 260,942 for the second-order birth regression sample and 40 out of 175,711 for the third-order birth regression sample. On average, there are 31 households per PSU in the regression samples.

Unobserved heterogeneity at the PSU level could cause variation in the manifestation of son preference. Early on in the data, this may include geographic differences in access to sex-determination technology. While ultrasound technology was widespread in India by the 1990s, access did continue expanding throughout that decade [Bhalotra and Cochrane (2010)]. Unfortunately, data collection on ultrasound usage is only available for a small subset of the sample and not at all for 1992–93, so it is not included in our regression models. We discuss the implications later in Section 4.2. However, PSU-level random effects (u_j in equation 1) can account for some of this unobserved heterogeneity in access to sex-determination technology.

There remains potential heterogeneity in social norms affecting the degree of son preference or at least the willingness to sex select. The preference for sons may itself differ across states or more locally between villages or throughout major urban areas

[13]This mirrors the NFHS sampling methodology which omits villages with fewer than five households from the survey.

https://doi.org/10.1017/dem.2020.5 Published online by Cambridge University Press

depending on economic and educational opportunities for women, religious traditions, and other social norms. While the control variables in X aim to account for such variation, they cannot capture all social norms or even wealth effects completely, so there may remain unobserved heterogeneity. Moreover, even for a particular degree of son preference, social norms may affect how individuals act upon that son preference, whether it manifests in sex-selective abortions, or even infanticide, or whether parents use the more natural "stopping rule", having children until the desired number of sons is born. Since responses to survey questions regarding sex selection cannot be considered reliable, in part due to the illegality of both sex-selective abortions and of infanticide, individual-level preferences cannot be directly measured [Arnold *et al.* (2002)]. Assuming social norms play a role in son preference and the proclivity to actually sex select, PSU-level random effects can account for some of this potential bias by controlling for unobserved community-level heterogeneity.

Sex selection at earlier births may affect the tendency to sex select at higher order births. If those having a third birth sex selected at previous births, they might be more likely to sex select again but to also have more sons, thereby biasing our results toward a weaker effect of having sons on the probability of having another son. If those having a third birth have not sex selected at previous births, they may be less willing and therefore less likely to sex select at the third order but also have more daughters, similarly biasing the estimated effects downward. Once again assuming social norms play a role in son preference and the proclivity to actually sex select, PSU-level random effects can account for some of this potential bias.

There could also be some sex selection occurring at first-order births, although past evidence concludes this an uncommon practice and is much more likely at higher order births [Ebenstein (2007), Portner (2016), Jha *et al.* (2011), Rosenblum (2013)].[14] However, to the extent that sex selection is occurring at the first order, this would bias the regression results that follow, which assume that the sex of the first-born child is exogenous, i.e., not affected by son preference. Again, PSU-level random effects can account for some of this potential bias. Moreover, analysis by Bhalotra and Cochrane (2010) concludes that unobserved mother-level heterogeneity in sex selection is not statistically significant in India.

4.2 Regression results

4.2.1 Sex selection at second-order births

Table 4 presents the results regressing probability that the second-born child is male on the sex of the first-born child and all of the control variables (not shown) for each sample year. Observations are weighted using sampling probability weights in these and all regressions that follow. If son preference is present and is manifesting in sex selection, then we expect a negative coefficient on the explanatory variable of interest, a first-born son. For each year, two models are presented: with and without average PSU children born per woman and its interaction with the sex of the first-born child. The latter investigates geographic variation in the fertility squeeze, that is, potentially more sex selection at lower order births where smaller families are the norm. Later, comparison with the regressions of third-order births will provide additional insight into the progression of the fertility squeeze over time. Coefficient

[14]Further drops in fertility could create a fertility squeeze at first-order births as in China where the one-child policy has had such an effect [Jha *et al.* (2011)].

https://doi.org/10.1017/dem.2020.5 Published online by Cambridge University Press

https://doi.org/10.1017/dem.2020.5 Published online by Cambridge University Press

Table 4. Random coefficients regressions: second-born male

	(1) 1992	(2) 1992	(3) 1998	(4) 1998	(5) 2005	(6) 2005	(7) 2015	(8) 2015
1st born male = 1	0.00858	−0.144	−0.0171	−0.143	−0.0613**	−0.117	−0.0796***	−0.354***
	(0.0187)	(0.0951)	(0.0188)	(0.0896)	(0.0205)	(0.0747)	(0.0224)	(0.0897)
Average PSU children born per woman		−0.0779***		−0.0438†		−0.0267		−0.0489
		(0.0223)		(0.0230)		(0.0287)		(0.0350)
1st born male = 1 × Average PSU children born per woman		0.0484		0.0415		0.0243		0.144***
		(0.0298)		(0.0287)		(0.0319)		(0.0437)
Var[1st born male (psu)]	0.0550***	0.0551***	0.0666***	0.0666***	0.145***	0.146***	0.453***	0.452***
	(0.0114)	(0.0114)	(0.0113)	(0.0113)	(0.0150)	(0.0150)	(0.0361)	(0.0361)
Var[cons (psu)]	0.0358***	0.0353***	0.0288***	0.0287***	0.0593***	0.0594***	0.190***	0.190***
	(0.00566)	(0.00564)	(0.00564)	(0.00563)	(0.00684)	(0.00684)	(0.0148)	(0.0148)
Observations	64,716	64,716	66,593	66,593	65,812	65,812	63,750	63,750
AIC	88,806.2	88,796.5	91,757.4	91,756.8	95,899.0	95,901.5	87,074.1	87,063.6
BIC	89,178.4	89,186.8	92,139.8	92,157.5	96,308.3	96,329.0	87,536.3	87,543.9

Observations are weighted using sampling probability weights. All regressions include state fixed effects and control variables. Clustered standard errors in parentheses.
†$p < 0.10$,
*$p < 0.05$,
**$p < 0.01$,
***$p < 0.001$.

estimates and statistical significance are similar across the logit, random-intercept, and random-coefficient models, but likelihood ratio tests favor the random-coefficient model over the basic logit and the random-intercept models in each case. Thus, only the random-coefficient estimates are discussed in the interest of brevity. The logit regression results are available in the Appendix.

Comparing results across years demonstrates little evidence of sex selection at second-order births in 1992 and 1998 but clear evidence in 2005 and 2015; where the first-born child is male, the likelihood is lower that the second-born will also be male, relative to the benchmark of a first-born daughter. The 2015 results indicate variation in the fertility squeeze across PSUs; where fertility is lower, there is an even lower probability that a first-born son is followed by the birth of another son. Notably, few of the control variables are statistically significant. Partner's education is associated with a lower probability of a son at second parity, but the results are not significant across all years. Respondents who live with their in-laws are more likely to sex select, although this too is not consistent across years. Respondents who live with their children (i.e., those reporting they are the mother of the household head) are more likely to sex select, suggesting that more traditional living arrangements whereby parents live with their married son are associated with sex selection. Generally, though, it seems that sex of the first-born child is the key predictor of whether a second-born child will be male, with stronger effects where the fertility transition is more advanced.

The magnitude of the random-coefficient estimates is sizable (0.15 in 2005 and 0.45 in 2015) relative to the estimated β coefficients. Unobserved heterogeneity at the PSU level is causing substantial variation in the manifestation of son preference at second-order births, even after controlling for PSU heterogeneity in fertility rates.

4.2.2 Sex selection at third-order births

The lack of evidence of sex selection for earlier years is likely due to the pace of demographic transition. India's TFR was approximately 3.8 in 1992–93, 3.4 in 1998–99, 2.9 in 2005–06, and 2.3 in 2015 (World Bank Indicators). With a higher fertility rate, the pressure to sex select may not manifest until higher order births. Table 5 presents the results regressing the probability of a third-born male birth on sex composition of the first- and second-born children. Again, if son preference is present and manifesting in sex selection, we expect a lower probability of the third-born child being male if at least one son is already born, and even more so if there are already two sons, relative to the benchmark of two previous female births. This effect is likely stronger where fertility rates are lower as families act to fulfill the dual desires of having sons but also a small family. Again, we present only the random-coefficient estimates as these are supported by likelihood ratio tests over the basic logit and the random-intercept models.

Unlike the regressions of second-order births, we see evidence of sex selection at the third-order birth even as early as 1992 once we account for variation in fertility rates. That is, the fertility squeeze was affecting third-order births at least as early as 1992, but not yet second-order births. The direct effect of having two sons is stronger than that of having one son and one daughter.[15] Geographic heterogeneity strengthens over time as

[15]It is possible that variation in sampling strategy or sample size across years is driving some of the regression results. A model that pools the data from all four waves and includes year interactions with the variables of interest yields similar results to those presented in Tables 4 and 5. The evidence of sex selection at second-order births appears more recently, in 2005 and 2015, while at third-order births,

https://doi.org/10.1017/dem.2020.5 Published online by Cambridge University Press

https://doi.org/10.1017/dem.2020.5 Published online by Cambridge University Press

Table 5. Random coefficients regressions: third-born male

	(1) 1992	(2) 1992	(3) 1998	(4) 1998	(5) 2005	(6) 2005	(7) 2015	(8) 2015
1st 2 born male = 1	0.0170	−0.528**	−0.0428	−0.583***	−0.0407	−0.358**	−0.198***	−1.061***
	(0.0315)	(0.165)	(0.0315)	(0.156)	(0.0373)	(0.136)	(0.0457)	(0.186)
1st 2 born—1 male and 1 female = 1	0.0189	−0.310*	−0.0266	−0.288*	−0.0272	−0.239*	−0.134***	−0.618***
	(0.0273)	(0.145)	(0.0275)	(0.139)	(0.0309)	(0.112)	(0.0359)	(0.150)
Average PSU children born per woman		−0.113**		−0.102**		−0.0834†		−0.200***
		(0.0372)		(0.0373)		(0.0429)		(0.0562)
1st 2 born male = 1 × Average PSU children born per woman		0.170***		0.173***		0.132*		0.420***
		(0.0506)		(0.0488)		(0.0552)		(0.0851)
1st 2 born—1 male and 1 female = 1 × Average PSU children born per woman		0.103*		0.0839†		0.0874†		0.239***
		(0.0445)		(0.0435)		(0.0465)		(0.0690)
Var[1st 2 born male (psu)]	0.264***	0.264***	0.249***	0.246***	0.528***	0.527***	1.666***	1.663***
	(0.0371)	(0.0371)	(0.0381)	(0.0378)	(0.0602)	(0.0601)	(0.169)	(0.168)
Var[1st 2 born—1 male/1 female (psu)]	0.0998***	0.100***	0.118***	0.118***	0.183***	0.182***	0.668***	0.665***
	(0.0170)	(0.0170)	(0.0188)	(0.0189)	(0.0222)	(0.0221)	(0.0698)	(0.0692)
Var[cons (psu)]	0.0216**	0.0211**	0.0177*	0.0174*	0.0747***	0.0747***	0.293***	0.293***
	(0.00769)	(0.00767)	(0.00789)	(0.00786)	(0.0114)	(0.0115)	(0.0302)	(0.0303)
Observations	48,087	48,087	47,536	47,536	42,744	42,744	37,304	37,304
AIC	66,658.4	66,647.5	65,778.7	65,767.9	64,210.0	64,206.6	46,273.6	46,247.7
BIC	67,036.0	67,051.4	66,164.6	66,180.1	64,617.2	64,639.7	46,725.5	46,725.2

Observations are weighted using sampling probability weights. All regressions include state fixed effects and control variables. Clustered standard errors in parentheses.
†$p < 0.10$,
*$p < 0.05$,
**$p < 0.01$,
***$p < 0.001$.

reflected by increasing magnitudes of the random-coefficient estimates (0.26 and 0.10 in 1992 up to 1.66 and 0.67 in 2015). As with the second-order birth regressions, this magnitude relative to the β coefficient estimates on sex composition indicates the importance of the unobserved, community-level heterogeneity. Again, few of the control variables are statistically significant. Respondent's education and living with one's children are associated with a higher probability of a son at third parity, but these are significant only for some years.

4.3 Robustness checks

We further examine the results at second- and third-order births by disaggregating the samples by rural–urban, regionally, and by respondent's age.

4.3.1 Urban–rural

Tables 6 and 7 segment the sample into respondents residing in rural vs urban areas. At second-order births (Table 6), there is weaker evidence of sex selection in urban areas in the 1990s that strengthens in statistical significance into the 2000s. For rural areas, sex selection at second-order births appears only in the 2000s. Evidence of geographic variation in the fertility squeeze is apparent only in the urban sub-sample. In other words, urban trends are the primary driver of the main results in Table 4. At third-order births (Table 7), there is evidence of sex selection as early as 1992 in both rural and urban areas, as in the main results in Table 5. Moreover, evidence of sex selection and variation in the fertility squeeze appear stronger in rural areas most recently and in urban areas in the 1990s. Comparing Tables 6 and 7, most recently, variation in the fertility squeeze emerges in urban areas at second-order births and strengthens in rural areas at third-order births, consistent with lower desired fertility in urban areas. Increasing magnitudes of the random-coefficient estimates also indicate growing heterogeneity in sex selection in both rural and urban areas.

4.3.2 By region

Geographic variation at the larger regional level is more closely investigated in Tables 8 and 9. Dyson and Moore (1983) demonstrated that a northwest–southeast divide exists in India regarding sex ratios and other demographic variables such as mortality and fertility, which is approximated by the contours of the Satpura hill range extending from the Gujarat–Maharashtra border on the west coast up to southern Bihar in the east where it joins the Chota Nagpur hills. States north of this demarcation have historically exhibited more masculine sex ratios than states south of the line, even pre-dating the spread of ultrasound technology. Accordingly, we now apply this classification by grouping observations into states falling in the regions "Northwest" and "Southeast".[16]

Table 8 indicates that the results for second-order births in Table 4 are being driven by the Northwest, such that sex selection appears to be occurring in the Northwest as

the evidence begins as early as 1992, with the probability of a third-born male decreasing in the number of sons born previously. These results are available upon request.

[16]Per Dyson and Moore, we place Bihar into the Northwest and West Bengal into the Southeast. Jharkhand, which was carved out of southern Bihar in 2000, fits with the Southeast grouping according to Dyson and Moore, who mapped more skewed sex ratios in northern Bihar and more balanced ones in southern Bihar.

https://doi.org/10.1017/dem.2020.5 Published online by Cambridge University Press

https://doi.org/10.1017/dem.2020.5 Published online by Cambridge University Press

Table 6. Random coefficients regressions: second-born male—rural vs urban

	(1) 1992	(2) 1992	(3) 1998	(4) 1998	(5) 2005	(6) 2005	(7) 2015	(8) 2015
Rural								
1st born male = 1	0.00795	−0.0892	−0.00595	−0.0826	−0.0494†	−0.0627	−0.0465†	−0.183†
	(0.0219)	(0.119)	(0.0224)	(0.120)	(0.0252)	(0.101)	(0.0248)	(0.101)
Average PSU children born per woman		−0.0703**		−0.0410		0.00987		−0.0494
		(0.0265)		(0.0287)		(0.0353)		(0.0381)
1st born male = 1 × Average PSU children born per woman		0.0300		0.0244		0.00543		0.0673
		(0.0365)		(0.0370)		(0.0400)		(0.0477)
Var[1st born male (psu)]	0.0570***	0.0565***	0.0758***	0.0759***	0.173***	0.172***	0.278***	0.278***
	(0.0129)	(0.0129)	(0.0138)	(0.0138)	(0.0198)	(0.0198)	(0.0327)	(0.0327)
Var[cons (psu)]	0.0338***	0.0333***	0.0307***	0.0306***	0.0669***	0.0670***	0.137***	0.137***
	(0.00618)	(0.00617)	(0.00654)	(0.00651)	(0.00864)	(0.00864)	(0.0149)	(0.0149)
Observations	44,808	44,808	46,373	46,373	37,794	37,794	45,758	45,758
AIC	65,740.3	65,736.2	68,201.1	68,202.5	67,070.7	67,074.5	58,636.1	58,637.3
BIC	66,088.7	66,102.0	68,559.6	68,578.5	67,446.5	67,467.3	59,063.9	59,082.6
Urban								
1st born male = 1	0.00899	−0.296†	−0.0494	−0.241†	−0.0894*	−0.204	−0.174***	−0:751***
	(0.0362)	(0.166)	(0.0342)	(0.143)	(0.0353)	(0.140)	(0.0463)	(0.194)

(*Continued*)

https://doi.org/10.1017/dem.2020.5 Published online by Cambridge University Press

Table 6. (*Continued.*)

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	1992	1992	1998	1998	2005	2005	2015	2015
Average PSU children born per woman		−0.0945*		−0.0266		−0.0849		−0.0441
		(0.0430)		(0.0408)		(0.0573)		(0.0850)
1st born male = 1 × Average PSU children born per woman		0.105†		0.0696		0.0593		0.348**
		(0.0557)		(0.0513)		(0.0717)		(0.109)
Var[1st born male (psu)]	0.0473*	0.0480*	0.0439*	0.0434*	0.0905***	0.0907***	0.932***	0.921***
	(0.0239)	(0.0239)	(0.0204)	(0.0204)	(0.0217)	(0.0219)	(0.105)	(0.105)
Var[cons (psu)]	0.0363**	0.0359**	0.0197	0.0199	0.0367***	0.0371***	0.320***	0.320***
	(0.0130)	(0.0129)	(0.0120)	(0.0120)	(0.0105)	(0.0105)	(0.0374)	(0.0373)
Observations	19,908	19,908	20,220	20,220	28,018	28,018	17,992	17,992
AIC	23,105.0	23,102.3	23,594.1	23,596.0	28,853.1	28,853.8	28,335.6	28,324.7
BIC	23,420.9	23,434.0	23,918.6	23,936.3	29,215.7	29,232.9	28,725.5	28,730.2

Observations are weighted using sampling probability weights. All regressions include state fixed effects and control variables. Clustered standard errors in parentheses.
†$p < 0.10$,
*$p < 0.05$,
**$p < 0.01$,
***$p < 0.001$.

https://doi.org/10.1017/dem.2020.5 Published online by Cambridge University Press

Table 7. Random coefficients regressions: third-born child male—rural vs urban

	(1) 1992	(2) 1992	(3) 1998	(4) 1998	(5) 2005	(6) 2005	(7) 2015	(8) 2015
Rural								
1st 2 born male = 1	0.0180	−0.404*	−0.0155	−0.359†	−0.0143	−0.352*	−0.164***	−1.108***
	(0.0365)	(0.199)	(0.0372)	(0.198)	(0.0452)	(0.176)	(0.0488)	(0.209)
1st 2 born—1 male and 1 female = 1	−0.00671	−0.294	−0.0000637	−0.131	−0.00563	−0.199	−0.113**	−0.558**
	(0.0316)	(0.179)	(0.0316)	(0.172)	(0.0369)	(0.144)	(0.0394)	(0.171)
Average PSU children born per woman		−0.0963*		−0.0526		−0.0791		−0.206**
		(0.0435)		(0.0444)		(0.0506)		(0.0627)
1st 2 born male = 1 × Average PSU children born per woman		0.129*		0.107†		0.132†		0.440***
		(0.0602)		(0.0604)		(0.0675)		(0.0930)
1st 2 born—1 male and 1 female = 1 × Average PSU children born per woman		0.0876		0.0410		0.0754		0.211**
		(0.0541)		(0.0526)		(0.0567)		(0.0765)
Var[1st 2 born male (psu)]	0.278***	0.279***	0.317***	0.315***	0.607***	0.606***	1.076***	1.077***
	(0.0418)	(0.0418)	(0.0478)	(0.0475)	(0.0767)	(0.0766)	(0.147)	(0.147)
Var[1st 2 born—1 male/1 female (psu)]	0.119***	0.120***	0.138***	0.138***	0.198***	0.198***	0.473***	0.470***
	(0.0203)	(0.0203)	(0.0228)	(0.0228)	(0.0268)	(0.0267)	(0.0625)	(0.0623)
Var[cons (psu)]	0.0141	0.0138	0.0191*	0.0188*	0.0842***	0.0843***	0.193***	0.192***
	(0.00822)	(0.00820)	(0.00935)	(0.00932)	(0.0141)	(0.0142)	(0.0306)	(0.0305)
Observations	34,405	34,405	34,654	34,654	26,395	26,395	28,344	28,344
AIC	50,725.1	50,723.5	50,603.5	50,605.3	47,351.6	47,351.7	34,076.0	34,055.3
BIC	51,079.9	51,103.5	50,967.0	50,994.2	47,727.9	47,752.6	34,496.9	34,501.0

(*Continued*)

https://doi.org/10.1017/dem.2020.5 Published online by Cambridge University Press

Table 7. (*Continued.*)

	(1) 1992	(2) 1992	(3) 1998	(4) 1998	(5) 2005	(6) 2005	(7) 2015	(8) 2015
Urban								
1st 2 born male = 1	0.0242	−0.842**	−0.135*	−1.052***	−0.105	−0.330	−0.308**	−0.798†
	(0.0621)	(0.298)	(0.0588)	(0.261)	(0.0657)	(0.256)	(0.114)	(0.454)
1st 2 born—1 male and 1 female = 1	0.102†	−0.512*	−0.114*	−0.559*	−0.0830	−0.281	−0.221**	−0.748*
	(0.0544)	(0.251)	(0.0557)	(0.247)	(0.0559)	(0.207)	(0.0830)	(0.346)
Average PSU children born per woman		−0.185*		−0.227**		−0.0880		−0.152
		(0.0738)		(0.0730)		(0.0943)		(0.138)
1st 2 born male = 1 × Average PSU children born per woman		0.287**		0.317***		0.110		0.272
		(0.0959)		(0.0890)		(0.122)		(0.236)
1st 2 born—1 male and 1 female = 1 × Average PSU children born per woman		0.204*		0.155†		0.0965		0.295
		(0.0809)		(0.0832)		(0.101)		(0.182)
Var[1st 2 born male (psu)]	0.196*	0.187*	0.0517	0.0472	0.337***	0.337***	4.259***	4.258***
	(0.0831)	(0.0831)	(0.0555)	(0.0536)	(0.0945)	(0.0948)	(0.685)	(0.685)
Var[1st 2 born—1 male/1 female (psu)]	0.0188	0.0182	0.0596	0.0590	0.137***	0.137***	1.508***	1.497***
	(0.0291)	(0.0292)	(0.0340)	(0.0340)	(0.0412)	(0.0411)	(0.255)	(0.251)
Var[cons (psu)]	0.0491*	0.0490*	0.0123	0.0117	0.0441*	0.0430*	0.672***	0.674***
	(0.0192)	(0.0193)	(0.0163)	(0.0162)	(0.0192)	(0.0192)	(0.0952)	(0.0954)
Observations	13,682	13,682	12,882	12,882	16,349	16,349	8,960	8,960
AIC	15,983.6	15,976.0	15,210.6	15,199.2	16,898.8	16,902.9	12,102.1	12,104.4
BIC	16,299.6	16,314.6	15,531.6	15,542.5	17,253.1	17,280.3	12,471.3	12,494.9

Observations are weighted using sampling probability weights. All regressions include state fixed effects and control variables. Clustered standard errors in parentheses. †$p < 0.10$, *$p < 0.05$, **$p < 0.01$, ***$p < 0.001$.

https://doi.org/10.1017/dem.2020.5 Published online by Cambridge University Press

Table 8. Random coefficients regressions: second-born male—by region

	(1) 1992	(2) 1992	(3) 1998	(4) 1998	(5) 2005	(6) 2005	(7) 2015	(8) 2015
Northwest								
1st born male = 1	−0.0189	−0.314*	−0.0175	−0.160	−0.133***	−0.353**	−0.123***	−0.774***
	(0.0259)	(0.142)	(0.0240)	(0.133)	(0.0300)	(0.110)	(0.0284)	(0.120)
Average PSU children born per woman		−0.0978**		−0.0421		−0.0667†		−0.195***
		(0.0300)		(0.0304)		(0.0395)		(0.0420)
1st born male = 1 × Average PSU children born per woman		0.0891*		0.0426		0.0848*		0.313***
		(0.0426)		(0.0398)		(0.0431)		(0.0547)
Var[1st born male (psu)]	0.0417**	0.0413**	0.0350*	0.0350*	0.149***	0.149***	0.196***	0.191***
	(0.0137)	(0.0138)	(0.0152)	(0.0152)	(0.0220)	(0.0219)	(0.0363)	(0.0360)
Var[cons (psu)]	0.0313***	0.0312***	0.0281**	0.0281**	0.0579***	0.0580***	0.0896***	0.0879***
	(0.00724)	(0.00723)	(0.00879)	(0.00880)	(0.0102)	(0.0102)	(0.0159)	(0.0160)
Observations	34,779	34,779	36,246	36,246	27,654	27,654	35,180	35,180
AIC	46,247.0	46,239.6	46,966.1	46,967.8	46,132.4	46,130.8	42,428.1	42,385.4
BIC	46,466.8	46,476.4	47,187.0	47,205.7	46,354.6	46,361.2	42,682.1	42,656.4
Southeast								
1st born male = 1	0.0384	−0.0748	−0.0176	−0.189	0.00284	−0.148	−0.0312	−0.0570
	(0.0273)	(0.135)	(0.0292)	(0.146)	(0.0280)	(0.112)	(0.0354)	(0.150)

(*Continued*)

https://doi.org/10.1017/dem.2020.5 Published online by Cambridge University Press

Table 8. (*Continued.*)

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	1992	1992	1998	1998	2005	2005	2015	2015
Average PSU children born per woman		−0.0743*		−0.0559		−0.0134		0.105
		(0.0343)		(0.0391)		(0.0441)		(0.0650)
1st born male = 1 × Average PSU children born per woman		0.0380		0.0629		0.0739		0.0151
		(0.0446)		(0.0530)		(0.0523)		(0.0814)
Var[1st born male (psu)]	0.0726***	0.0725***	0.0955***	0.0953***	0.138***	0.136***	0.796***	0.794***
	(0.0190)	(0.0191)	(0.0175)	(0.0175)	(0.0202)	(0.0202)	(0.0733)	(0.0731)
Var[cons (psu)]	0.0405***	0.0396***	0.0315***	0.0313***	0.0611***	0.0612***	0.313***	0.311***
	(0.00888)	(0.00881)	(0.00779)	(0.00773)	(0.00917)	(0.00915)	(0.0282)	(0.0280)
Observations	29,937	29,937	30,347	30,347	38,158	38,158	28,487	28,487
AIC	42,559.0	42,557.2	44,794.7	44,795.7	49,755.0	49,757.4	44,465.0	44,463.9
BIC	42,816.5	42,831.3	45,060.9	45,078.6	50,045.7	50,056.6	44,762.3	44,777.7

Observations are weighted using sampling probability weights. All regressions include state fixed effects and control variables. Clustered standard errors in parentheses.
†$p < 0.10$,
*$p < 0.05$,
**$p < 0.01$,
***$p < 0.001$.

https://doi.org/10.1017/dem.2020.5 Published online by Cambridge University Press

Table 9. Random coefficients regressions: third-born male—by region

	(1) 1992	(2) 1992	(3) 1998	(4) 1998	(5) 2005	(6) 2005	(7) 2015	(8) 2015
Northwest								
1st 2 born male = 1	0.0507	−0.574*	−0.00202	−0.826***	−0.0987†	−0.704***	−0.152**	−1.015***
	(0.0420)	(0.242)	(0.0391)	(0.220)	(0.0528)	(0.203)	(0.0494)	(0.216)
1st 2 born—1 male and 1 female = 1	0.0213	−0.407†	−0.0229	−0.546**	−0.0501	−0.422*	−0.0860*	−0.536**
	(0.0368)	(0.212)	(0.0343)	(0.198)	(0.0450)	(0.169)	(0.0403)	(0.173)
Average PSU children born per woman		−0.154**		−0.137**		−0.136*		−0.187**
		(0.0517)		(0.0491)		(0.0585)		(0.0636)
1st 2 born male = 1 × Average PSU children born per woman		0.186**		0.243***		0.225**		0.393***
		(0.0714)		(0.0653)		(0.0747)		(0.0957)
1st 2 born—1 male and 1 female = 1 × Average PSU children born per woman		0.127*		0.155**		0.139*		0.208**
		(0.0630)		(0.0590)		(0.0637)		(0.0775)
Var[1st 2 born male (psu)]	0.179***	0.181***	0.130**	0.127**	0.511***	0.507***	0.695***	0.694***
	(0.0404)	(0.0404)	(0.0400)	(0.0398)	(0.0810)	(0.0805)	(0.124)	(0.123)
Var[1st 2 born—1 male/1female (psu)]	0.0808***	0.0812***	0.0662**	0.0663**	0.206***	0.205***	0.265***	0.264***
	(0.0210)	(0.0209)	(0.0215)	(0.0215)	(0.0311)	(0.0309)	(0.0521)	(0.0519)
Var[cons (psu)]	0.0119	0.0107	0.0110	0.0107	0.0857***	0.0857***	0.0841***	0.0836***
	(0.00942)	(0.00933)	(0.00974)	(0.00975)	(0.0160)	(0.0160)	(0.0240)	(0.0240)
Observations	26696	26,696	27,421	27,421	19,579	19,579	22,532	22,532
AIC	36,228.4	36,220.8	36,413.5	36,402.2	34,640.2	34,632.3	27,371.2	27,357.0
BIC	36,457.7	36,474.8	36,643.6	36,656.9	34,868.7	34,884.6	27,627.9	27,637.8

(*Continued*)

https://doi.org/10.1017/dem.2020.5 Published online by Cambridge University Press

Table 9. (*Continued.*)

	(1) 1992	(2) 1992	(3) 1998	(4) 1998	(5) 2005	(6) 2005	(7) 2015	(8) 2015
Southeast								
1st 2 born male = 1	−0.0193	−0.464†	−0.0905†	−0.287	0.0227	−0.346	−0.267**	−1.195**
	(0.0474)	(0.238)	(0.0516)	(0.260)	(0.0527)	(0.222)	(0.0932)	(0.364)
1st 2 born—1 male and 1 female = 1	0.0168	−0.244	−0.0296	−0.0725	−0.00285	−0.226	−0.201**	−0.738*
	(0.0408)	(0.208)	(0.0443)	(0.223)	(0.0420)	(0.166)	(0.0702)	(0.292)
Average PSU children born per woman		−0.0691		−0.0646		−0.0789		−0.220†
		(0.0566)		(0.0658)		(0.0679)		(0.117)
1st 2 born male = 1 × Average PSU children born per woman		0.147†		0.0699		0.173		0.497**
		(0.0769)		(0.0906)		(0.102)		(0.181)
1st 2 born—1 male and 1 female = 1 × Average PSU children born per woman		0.0860		0.0152		0.104		0.290*
		(0.0669)		(0.0775)		(0.0760)		(0.147)
Var[1st 2 born male (psu)]	0.387***	0.386***	0.429***	0.428***	0.543***	0.544***	4.171***	4.159***
	(0.0670)	(0.0670)	(0.0767)	(0.0767)	(0.0889)	(0.0888)	(0.544)	(0.543)
Var[1st 2 born—1 male/1 female (psu)]	0.132***	0.132***	0.189***	0.188***	0.157***	0.157***	1.920***	1.912***
	(0.0283)	(0.0282)	(0.0344)	(0.0344)	(0.0331)	(0.0330)	(0.234)	(0.232)
Var[cons (psu)]	0.0373**	0.0373**	0.0320*	0.0313*	0.0617***	0.0615***	0.857***	0.861***
	(0.0129)	(0.0129)	(0.0137)	(0.0136)	(0.0168)	(0.0168)	(0.0910)	(0.0913)
Observations	21,391	21,391	20,115	20,115	23,165	23,165	14,731	14,731
AIC	30,414.9	30,416.3	29,346.1	29,350.0	29,553.0	29,554.6	18,517.0	18,513.8
BIC	30,678.0	30,703.2	29,615.0	29,642.6	29,842.8	29,868.5	18,805.7	18,825.3

Observations are weighted using sampling probability weights. All regressions include state fixed effects and control variables. Clustered standard errors in parentheses.
†$p < 0.10$,
*$p < 0.05$,
**$p < 0.01$,
***$p < 0.001$.

early as 1992, and certainly in the 2000s, but not in the rest of the country. PSU-level heterogeneity is significant and generally growing over time in both regions but especially in the Southeast. Table 9 indicates that the results for third-order births in Table 5 are also being driven primarily by the Northwest until most recently, but by 2015, evidence of sex selection emerges in the Southeast also. The fertility squeeze is finally affecting even this region of the country where sex ratios have traditionally remained more balanced. Again, PSU-level heterogeneity is growing over time in both regions, and the random-coefficient estimates are particularly large for the Southeast.

4.3.3 By respondent age

The regressions pool women at various stages of fertility, which may mask changing social norms. Second- or third-order births of older women could appear in numerous regression years. Tables 10 and 11 segment the sample into respondents less than 40 years old and those 40 and older to investigate whether sex selection is indeed occurring recently or whether the results in Tables 4 and 5 are being driven by remnants of sex-selective practices of older cohorts whose fertility is for the most part completed. Results are similar for small variations in the age groupings. For the older cohort, the PSU average number of children born to respondents is re-calculated to include only the birth histories of women 40 years and older; the fertility behavior of younger women would not have affected past fertility choices of older women, although underlying trends impacting the fertility of younger women would evolve over a longer time frame. Consistent with the main results in Tables 4 and 5, there is evidence of sex selection at second-order births in the 2000s (Table 10) and at least some at third-order births as early as 1992 (Table 11) across younger and older women. Most recently, in 2015, there is variation in the fertility squeeze among both age groups for both second- and third-order births (column 8). Among the older women, these births likely occurred several years in the past and therefore represent the fertility squeeze with some lag. Weaker statistical significance among younger women in column 8 of Table 11 suggests that variation in fertility norms has less effect at third-order births now that India's TFR is below 3. Taken together, Tables 10 and 11 demonstrate sex selection among women earlier in their childbearing years up until most recently, confirming that the practice is not a fading phenomenon.

4.3.4 Caste

As a robustness check, all of the above regressions are repeated controlling for caste. Specifically, the NFHS identifies each respondent by whether she belongs to a scheduled caste (19–22%), scheduled tribe (9–10%), or other backwards class (33–45%). The 1992–93 regressions are omitted because of lack of data for this variable. Regression results are not importantly altered by this addition and are available upon request.

4.3.5 Ultrasound access

Because of data limitations, none of the regression results presented account for access to sex-determination technology, and this could be driving some of the outcomes. As the access to ultrasound technology spread, so would evidence of sex selection, as in Table 4 where sex selection at second-order births shows up only more recently in the 2000s. However, evidence of sex selection at third-order births as early as 1992

https://doi.org/10.1017/dem.2020.5 Published online by Cambridge University Press

https://doi.org/10.1017/dem.2020.5 Published online by Cambridge University Press

Table 10. Random coefficients regressions: second-born male—by age group

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	1992	1992	1998	1998	2005	2005	2015	2015
Ages 15–39								
1st born male = 1	−0.00534	−0.135	−0.0355	−0.135	−0.0914***	−0.232*	−0.0927**	−0.361**
	(0.0215)	(0.111)	(0.0218)	(0.102)	(0.0249)	(0.0908)	(0.0285)	(0.115)
Average PSU children born per woman		−0.0656*		−0.0284		−0.0586		−0.105*
		(0.0264)		(0.0261)		(0.0333)		(0.0440)
1st born male = 1 × Average PSU children born per woman		0.0412		0.0324		0.0600		0.138*
		(0.0348)		(0.0327)		(0.0385)		(0.0556)
Var[1st born male (psu)]	0.0629***	0.0628***	0.0811***	0.0812***	0.229***	0.229***	0.655***	0.656***
	(0.0163)	(0.0163)	(0.0152)	(0.0152)	(0.0245)	(0.0245)	(0.0603)	(0.0603)
Var[con (psu)]	0.0559***	0.0556***	0.0395***	0.0396***	0.0928***	0.0930***	0.333***	0.332***
	(0.00812)	(0.00810)	(0.00790)	(0.00789)	(0.0102)	(0.0102)	(0.0256)	(0.0256)
Observations	47,635	47,635	48,700	48,700	45,806	45,806	41,085	41,085
AIC	66,345.2	66,342.1	67,826.6	67,829.0	67,886.9	67,885.7	55,589.3	55,583.6
BIC	66,704.8	66,719.3	68,196.0	68,215.9	68,279.9	68,296.1	56,029.0	56,040.6
Ages 40–49								
1st born male = 1	0.0594	−0.278?	0.0379	−0.180	0.0178	−0.135	−0.0518	−0.498***
	(0.0390)	(0.154)	(0.0369)	(0.135)	(0.0401)	(0.126)	(0.0441)	(0.147)
		−0.103***		−0.0988***		−0.0846**		−0.248***

https://doi.org/10.1017/dem.2020.5 Published online by Cambridge University Press

Average PSU children born per woman, ages 40+								
		(0.0225)		(0.0236)		(0.0260)		(0.0297)
1st born male = 1 × Average PSU children born per woman, ages 40+		0.0642*		0.0452†		0.0349		0.120**
		(0.0294)		(0.0273)		(0.0286)		(0.0372)
Var[1st born male (psu)]	0.230***	0.235***	0.261***	0.262***	0.553***	0.558***	1.831***	1.829***
	(0.0542)	(0.0543)	(0.0537)	(0.0537)	(0.0739)	(0.0741)	(0.168)	(0.168)
Var[cons (psu)]	0.146***	0.142***	0.118***	0.112***	0.252***	0.247***	0.723***	0.710***
	(0.0282)	(0.0280)	(0.0250)	(0.0248)	(0.0309)	(0.0308)	(0.0611)	(0.0606)
Observations	17,081	17,081	17,893	17,893	20,006	20,006	22,665	22,665
AIC	22,316.4	22,295.9	23,821.1	23,802.6	27,494.1	27,483.5	30,164.6	30,086.8
BIC	22,634.0	22,629.0	24,148.4	24,145.5	27,849.8	27,855.0	30,574.1	30,512.3

Observations are weighted using sampling probability weights. All regressions include state fixed effects and control variables. Clustered standard errors in parentheses.
†$p < 0.10$,
*$p < 0.05$,
**$p < 0.01$,
***$p < 0.001$.

https://doi.org/10.1017/dem.2020.5 Published online by Cambridge University Press

Table 11. Random coefficients regressions: third-born male—by age group

	(1) 1992	(2) 1992	(3) 1998	(4) 1998	(5) 2005	(6) 2005	(7) 2015	(8) 2015
Ages 15–39								
1st 2 born male = 1	−0.0125	−0.467*	−0.0499	−0.383*	−0.106*	−0.376*	−0.137*	−0.574*
	(0.0388)	(0.182)	(0.0384)	(0.182)	(0.0486)	(0.174)	(0.0661)	(0.266)
1st 2 born—1 male and 1 female = 1	0.0122	−0.0594	−0.0331	−0.248*	−0.0450	−0.159	−0.0849†	−0.352*
	(0.0330)	(0.114)	(0.0325)	(0.115)	(0.0391)	(0.123)	(0.0486)	(0.158)
Average PSU children born per woman		−0.0823**		−0.0301		−0.0494		−0.0985
		(0.0316)		(0.0348)		(0.0446)		(0.0612)
1st 2 born male = 1 × Average PSU children born per woman		0.140*		0.105†		0.110		0.216†
		(0.0549)		(0.0554)		(0.0683)		(0.119)
1st 2 born—1 male and 1 female = 1 × Average PSU children born per woman, ages 40+		0.0290†		0.0122		0.00273		0.0145
		(0.0158)		(0.0159)		(0.0211)		(0.0294)
Var[1st 2 born male (psu)]	0.432***	0.433***	0.424***	0.418***	0.985***	1.003***	2.703***	2.548***
	(0.0601)	(0.0606)	(0.0651)	(0.0653)	(0.120)	(0.122)	(0.346)	(0.323)
Var[1st 2 born—1 male/1 female (psu)]	0.152***	0.154***	0.165***	0.167***	0.329***	0.327***	1.010***	1.005***
	(0.0269)	(0.0273)	(0.0286)	(0.0289)	(0.0413)	(0.0415)	(0.122)	(0.122)
Var[cons (psu)]	0.0292*	0.0299*	0.0266*	0.0267*	0.132***	0.130***	0.504***	0.499***
	(0.0118)	(0.0119)	(0.0122)	(0.0123)	(0.0181)	(0.0182)	(0.0565)	(0.0566)
Observations	32,857	32,249	32,164	31,697	27,146	26,787	20,911	20,718
AIC	46,394.0	45,796.7	44,976.3	44,417.4	41,652.2	41,317.6	25,435.0	25,177.4
BIC	46,755.2	46,190.6	45,344.9	44,818.8	42,038.0	41,735.6	25,856.3	25,629.9

https://doi.org/10.1017/dem.2020.5 Published online by Cambridge University Press

Ages 40–49								
1st 2 born male = 1	0.111†	−0.281	−0.00679	−0.866***	0.103	−0.580*	−0.260***	−1.485***
	(0.0631)	(0.259)	(0.0574)	(0.217)	(0.0699)	(0.237)	(0.0781)	(0.281)
1st 2 born—1 male and 1 female = 1	0.0474	−0.127	−0.0148	−0.244	0.0383	−0.397*	−0.186**	−0.851***
	(0.0506)	(0.208)	(0.0496)	(0.194)	(0.0572)	(0.197)	(0.0618)	(0.226)
Average PSU children born per woman, ages 40+		−0.0655*		−0.0802*		−0.153***		−0.296***
		(0.0320)		(0.0335)		(0.0378)		(0.0445)
1st 2 born male = 1 × Average PSU children born per woman, ages 40+		0.0729		0.173***		0.150**		0.301***
		(0.0482)		(0.0429)		(0.0516)		(0.0653)
1st 2 born—1 male and 1 female = 1 × Average PSU children born per woman, ages 40+		0.0318		0.0466		0.0951*		0.162**
		(0.0382)		(0.0384)		(0.0432)		(0.0522)
Var[1st 2 born male (psu)]	1.199***	1.200***	0.600***	0.573***	1.946***	1.946***	3.311***	3.292***
	(0.182)	(0.182)	(0.122)	(0.119)	(0.279)	(0.279)	(0.438)	(0.434)
Var[1st 2 born—1 male/1 female (psu)]	0.417***	0.418***	0.328***	0.328***	0.764***	0.763***	1.984***	1.953***
	(0.0790)	(0.0791)	(0.0652)	(0.0652)	(0.105)	(0.105)	(0.211)	(0.209)
Var[cons (psu)]	0.108**	0.105**	0.0558*	0.0543*	0.265***	0.257***	0.800***	0.785***
	(0.0330)	(0.0331)	(0.0268)	(0.0266)	(0.0416)	(0.0413)	(0.0865)	(0.0850)
Observations	15,230	15,230	15,372	15,372	15,598	15,598	16,393	16,393
AIC	19,998.0	19,997.5	20,686.8	20,671.6	21,875.4	21,855.1	19,980.1	19,923.4
BIC	20,326.1	20,348.5	21,023.0	21,030.7	22,235.2	22,237.8	20,388.4	20,354.9

Observations are weighted using sampling probability weights. All regressions include state fixed effects and control variables. Clustered standard errors in parentheses.
†$p < 0.10$,
*$p < 0.05$,
**$p < 0.01$,
***$p < 0.001$.

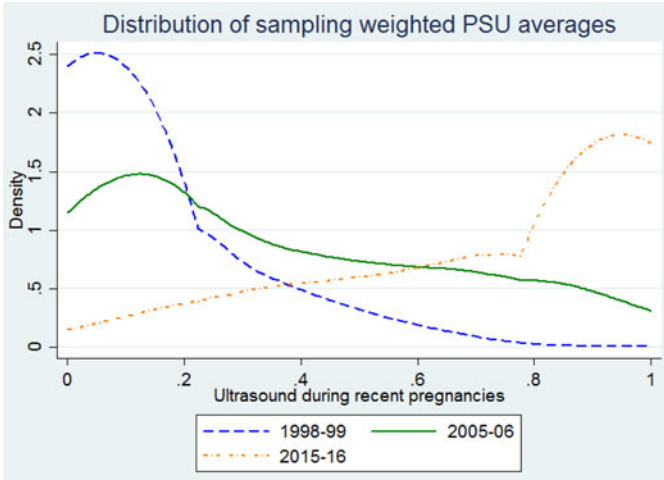


Figure 3. Proportion of respondents reporting at least one ultrasound, conditional on having been pregnant within the past 5 years, by survey wave, NFHS.
Note: First PSU averages are calculated, then probability sampling weights are applied when generating the density distributions. All available observations are included in the graphs, not just those used in the regression samples. The sample sizes here are: 39,156 for 1998–99, 206,048 for 2005–06, and 190,959 for 2015–16.

(Table 5) suggests that access to technology is not driving these results. Variation in access to sex-determination technology might also explain the PSU-level heterogeneity indicated by the random-coefficient estimates. Figure 3 illustrates the rising prevalence of ultrasound usage over time. Variation increased between 1998 and 2005 and then decreased again by 2015 as the distribution shifted strongly to the right. Thus, it is unlikely that variation in ultrasound usage explains the strong, monotonic increase in unobserved heterogeneity indicated by the mixed-effects regression results. Nevertheless, data limitations make it difficult to properly assess the role of technology diffusion in driving the heterogeneity results.

5. Conclusion

The tendency for son preference to decline with economic development may alleviate concerns about India's masculinized SRB [Bongaarts (2013)]. Yet falling desired family size can have the perverse effect of intensifying these imbalances, even as actual son preference weakens, by pressuring parents to sex select at earlier parities. Indeed, the analysis presented here demonstrates sex selection at increasingly earlier parities, consistent with the "intensification effect" described by Das Gupta and Bhat (1997).

This paper is the first to combine all four waves of India's NFHS to examine the influence of the fertility transition on the country's notoriously skewed SRB. The regression results demonstrate a fertility squeeze at third-order births in even the earliest survey years (1992–93), while the effect at second-order births does not emerge until more recently, beginning with the 2005–06 survey and strengthening for 2015–16. Accordingly, India's TFR fell from 3.8 to 2.35 during this time [World Bank Indicators (2015)]. The analysis also accounts for unobserved heterogeneity at

https://doi.org/10.1017/dem.2020.5 Published online by Cambridge University Press

the village or urban neighborhood level that may affect the degree of son preference and its propensity to manifest in sex selection. Unlike Bhalotra and Cochrane (2010) who conclude that heterogeneity at the individual mother level is of little importance in India, we show that heterogeneity at the community level is considerable in magnitude and has increased over time, even after controlling for geographic variation in the pace of fertility decline. Future research may determine the drivers of this heterogeneity, whether they are access to health care facilities willing to provide abortions or specific cultural or socioeconomic factors not captured in the estimated models. Without explicitly controlling for ultrasound access, it is unclear to what degree technology access is driving the unobserved heterogeneity. Having controlled for desired fertility, we find no systematic impact of religion, caste, household structure, exposure to media, or education or wealth levels on the propensity to sex select in the regression analysis here, supporting Basu's argument of cultural convergence throughout India as lower socioeconomic groups imitate upper ones, which tend to have both stronger son preference and lower fertility [Basu (1999)].[17] Future research may further examine cultural and socioeconomic factors by including more refined measures of caste, access to sex determination technology, and economic status.

As it has become easier for parents to control the sex composition of their children, the health and education outcomes of the so-called "wanted" daughters have been found to improve [Mishra et al. (2004), Chamarbagwala (2011), Hu and Schlosser (2012), Rosenblum (2013), Hazan and Zoabi (2015)], making sex-selective abortion a mechanism by which discrimination against girls is being replaced by discrimination against female fetuses [Anukriti *et al.* (2018)]. Conditional cash transfer programs that subsidize daughters have met with mixed success and are, of course, subject to funding availability [Anukriti (2018)]. Future research can pinpoint specific villages and neighborhoods that are more prone to the practice of sex-selective abortion. It is possible that social network effects play a role in the local geographic heterogeneity identified by the regression analysis here. In that case, policies may be able to exploit these social network effects to better target those couples most likely to sex select.

As desired fertility continues to decline, even decreased rates of sex selection could exacerbate sex ratio imbalances because the number of missing girls weighs disproportionately on the SRB of a small birth cohort. Accordingly, our results demonstrate sex selection at increasingly lower order births. India's SRB could eventually re-balance, as was the case in South Korea, although likely not until well after fertility rates have stabilized and further economic development counters the fertility squeeze effect. On the other hand, India may be on a path of regional convergence toward a permanently skewed SRB if the willingness to sex select persists at very low fertility [Basu (1999), Diamond-Smith and Bishai (2015), Rajan *et al.* (2017)]. Our results confirm the Northwest is driving much of the aggregate trends in sex selection. But they also reveal evidence of sex selection most recently, in 2015, at third-order births in the Southeast where sex ratios have historically

[17]Partner's education is associated with a lower probability of a son at second parity, while education of the respondent is associated with a higher probability at third parity, but the results are not significant across all years. Echavarri and Ezcurra (2010) find a non-linear effect of education on sex selection in India, which may explain the inconsistency across our results. In some regressions, respondents who live with their in-laws or are the mother of the household head are more likely to sex select, confirming the influence of tradition on son preference, although this too is not consistent across years.

https://doi.org/10.1017/dem.2020.5 Published online by Cambridge University Press

remained more balanced, validating anecdotal evidence of widespread discrimination against baby girls throughout the South [Madan and Breuning (2014)]. The increase in unobserved heterogeneity in sex selection is also particularly strong for the Southeast. This might reflect a cultural shift in that region, but it could also be indicative of a different threshold for the fertility squeeze, with underlying son preference only recently manifesting is skewed sex ratios in this region. Kashyap and Villavicencio (2016) describe the preconditions for couples to sex select as RWA, or ready (fertility squeeze), willing (son preference), and able (technology). Couples may have long been "willing" to sex select but have only recently been "ready" to do so as desired family size falls further. With fertility falling, sex selection occurring at increasingly earlier parities, and sex selection emerging outside the Northwest, India's SRB is poised to continue rising in the coming years.

References

Anukriti, S (2018) Financial incentives and the fertility-sex ratio trade-off. *American Economic Journal: Applied Economics* 10(2), 27–57.

Anukriti, S, Sonia Bhalotra and Hiu Tam (2018) On the quantity and quality of girls: fertility, parental investments, and mortality. Boston College Working Papers in Economics.

Arnold, Fred, Minja Kim Choe and T.K. Roy (1998) Son preference, the family-building process and child mortality in India. *Population Studies* 52(3), 301–315 (November).

Arnold, Fred, Sunita Kishor and T.K. Roy (2002) Sex-selective abortions in India. *Population and Development Review* 28(4), 759–785 (December).

Arokiasamy, Perianayagam (2007) Sex ratio at birth and excess female child mortality in India: trends, differentials and regional patterns. In Isabelle Attané and Christophe Z. Guilmoto (eds.), *Watering the Neighbour's Garden: The Growing Female Deficit in Asia*, pp. 49–72. Paris: CICRED.

Balbo, Nicoletta and Nicola Barban (2014) Does fertility behavior spread among friends? *American Sociological Review* 79(3), 412–431.

Basu, Alaka Malwade (1999) Fertility decline and increasing gender imbalance in India, including a possible South Indian turnaround. *Development and Change* 30, 237–263.

Bernardi, Laura and Andreas Klärner (2014) Social networks and fertility. *Demographic Research* 30(22), 641–670.

Bhalotra, Sonia R. and Tom Cochrane (2010) Where have all the young girls gone? Identification of sex selection in India. IZA Discussion Paper No. 5381 (December).

Bhaskar V. (2011) Sex selection and gender balance. *American Economic Journal, Microeconomics* 3, 214–244.

Bhat, P.N. Mari and A.J. Francis Zavier (2003) Fertility decline and gender bias in northern India. *Demography* 40(4), 637–657.

Bloom, David E., David Canning, Isabel Günther and Sebastian Linnemayer (2008) Social interactions and fertility in developing countries. Program on the Global Demography of Aging Working Paper No. 34.

Bongaarts, John (2013) The implementation of preferences for male offspring. *Population and Development Review* 39(2), 185–208 (June).

Borker, Girija, Jan Eeckhout, Nancy Luke, Shantidani Minz, Kaivan Munshi and Soumya Swaminathan (2018) Wealth, marriage, and sex selection. Working Paper.

Census India. SRS Statistical Report 2015. Chapter 3, pp. 49-119. Accessed February 22, 2018, http://censusindia.gov.in/vital_statistics/SRS_Reports_2015.html.

Chamarbagwala, Rubiana (2011) Sibling composition and selective gender-based survival bias. *Journal of Population Economics* 24, 935–955.

Chung, Woojin and Monica Das Gupta (2007) The decline of son preference in South Korea: the roles of development and public policy. *Population and Development Review* 33(4), 757–783.

Clark, Shelley (2000) Son preference and sex composition of children: evidence from India. *Demography* 37 (1), 95–108 (February).

Croll, Elisabeth J. (2000). *Endangered Daughters: Discrimination and Development in Asia*. London: Routledge.

https://doi.org/10.1017/dem.2020.5 Published online by Cambridge University Press

Croll, Elisabeth J. (2002) Fertility decline, family size and female discrimination: a study of reproductive management in East and South Asia. *Asia-Pacific Population Journal* 17(2), 11–38 (June).

Dandona, Lalit and India State-Level Disease Burden Initiative Collaborators (2017) Nations within a nation: variations in epidemiological transition across the states of India, 1990–2016 in the Global Burden of Disease Study. *The Lancet* 390(10111), 2437–2460 (December).

Das Gupta, Monica (2005) Explaining Asia's 'Missing women': a new look at the data. *Population and Development Review* 31(3), 529–535 (September).

Das Gupta, Monica and P.N. Mari Bhat (1997) Fertility decline and increased manifestation of sex bias in India. *Population Studies* 51(3), 307–315.

Dharmalingam, Arunachalam, Sowmya Rajan and S. Philip Morgan (2014) The determinants of low fertility in India. *Demography* 51, 1451–1475.

Diamond-Smith, Nadia and David Bishai (2015) Evidence of self-correction of child sex ratios in India: a district-level analysis of child sex ratios from 1981 to 2011. *Demography* 52, 641–666.

Dubuc, Sylvie and Devinderjit S. Sivia (2017) Gender preferences and fertility effects on sex-composition: linking behaviour and macro-level effects. Working Paper.

Dyson, Tim and Mick Moore (1983) On kinship structure, female autonomy, and demographic behavior in India. *Population and Development Review* 9(1), 35–60 (March).

Ebenstein, Avraham Y. (2007) Fertility choices and sex selection in Asia: analysis and policy, Mimeo.

Echávarri, Rebeca A. and Roberto Ezcurra (2010) Education and gender bias in the sex ratio at birth: evidence from India. *Demography* 47(1), 249–268 (February).

Gu, Baochang and Krishna Roy (1995) Sex ratio at birth in China, with reference to other areas in East Asia: what we know. *Asia Pacific Population Journal* 10(3), 17–42.

Guilmoto, Christophe Z. (2009) The sex ratio transition in Asia. *Population and Development Review* 35(3), 519–549.

Guilmoto, Christophe Z. (2012) Sex imbalances at birth: Current trends, consequences and policy implications. UNFPA Asia and the Pacific Regional Office.

Hazan, Moshe and Hosny Zoabi (2015) Sons or daughters? Sex preferences and the reversal of the gender educational gap. *Journal of Demographic Economics* 81(2), 179–201.

Hesketh, Therese, Li Lu and Zhu Wei Xing (2011) The consequences of son preference and sex-selective abortion in China and other Asian countries. *Canadian Medical Association Journal* 183(12), 1374–1377 (September).

Hu, Luojia and Analía Schlosser (2012) Trends in prenatal sex selection and girls' nutritional status in India. *CESifo Economic Studies* 58(2), 348–372.

Jayachandran, Seema (2015) The roots of gender inequality in developing countries. *Annual Review of Economics* 7, 63–88.

Jayachandran, Seema (2017) Fertility decline and missing women. *American Economic Journal, Applied Economics* 9(1), 118–139.

Jennings, Julia A., Allison R. Sullivan and J. David Hacker (2012) Intergenerational transmission of reproductive behavior during the demographic transition. *Journal of Interdisciplinary History* 42(4), 543–569.

Jha, Prabhat, Maya A. Kesler, Rajesh Kumar, Faujdar Ram, Usha Ram, Lukasz Aleksandrowicz, Diego G. Bassani, Shailaja Chandra and Jayant K. Banthia (2011) Trends in selective abortions of girls in India: analysis of nationally representative birth histories from 1990 to 2005 and census data from 1991 to 2011. *Lancet (London, England)* 377, 1921–1928.

Kashyap, Ridhi and Francisco Villavicencio (2016) The dynamics of son preference, technology diffusion, and fertility decline underlying distorted sex ratios at birth: a simulation approach. *Demography* 53, 1261–1281.

Kashyap, Ridhi and Francisco Villavicencio (2017) An agent-based model of sex ratio at birth distortions. In André Grow and Jan Van Bavel (eds.), *Agent-Based Modelling in Population Studies: Concepts, Methods and Applications*, pp. 343–367. The Springer Series on Demographic Methods and Population Analysis, vol. 41. Cham: Springer.

Kolk, Martin, Daniel Cownden and Magnus Enguist (2014) Correlations in fertility across generations: Can low fertility persist? *Proceedings of the Royal Society B* 281, 20132561.

Li, Hongbin, Junjian Yi and Junsen Zhang (2011) Estimating the effect of the one-child policy on the sex ratio imbalance in China. *Demography* 48, 1535–1557.

https://doi.org/10.1017/dem.2020.5 Published online by Cambridge University Press

Madan, Kamlesh and Martijn H. Breuning (2014) Impact of prenatal technologies on the sex ratio in India: an overview. *Genetics in Medicine* 16(6), 425–432 (June).
Madhavan, Sangeetha, Alayne Adams and Dominique Simon (2003) Women's networks and the social world of fertility behavior. *International Family Planning Perspectives* 29(2), 58–68.
Mayer, Peter (1999) India's falling sex ratios. *Population and Development Review* 25(2), 323–343 (June).
Mishra, Vinod, T.K. Roy and Robert D. Retherford (2004) Sex differentials in childhood feeding, health care, and nutritional status in India. *Population and Development Review* 30(2), 269–295.
Munshi, Kaivan and Jacques Myaux (2006) Social norms and the fertility transition. *Journal of Development Economics* 80, 1–38.
Patel, Archana B., Neetu Badhoniya, Manju Mamtani and Hemant Kulkarni (2013) Skewed sex ratios in India, 'Physician, heal thyself'. *Demography* 50, 1129–1134.
Portner, Claus C. (2016) Sex selective abortions, fertility and birth spacing. *World Bank Policy Research Working Paper* No. 7189.
Rajan, S. Irudaya, Sharada Srinivasan and Arjun S. Bedi (2017) Update on trends in sex ratio at birth in India. *Economic and Political Weekly* 52(11), 14–16 (March).
Robitaille, Marie-Claire (2013) Determinants of stated son preference in India: are men and women different? *Journal of Development Studies* 49(5), 657–669.
Rosenblum, Daniel (2013) The effect of fertility decisions on excess female mortality in India. *Journal of Population Economics* 26(1), 147–180.
Sen, Amartya (1990) More than 100 million women are missing. *New York Review of Books* 37(20), 61–66 (December).
World Bank. Measure DHS: Demographic and Health Surveys. http://microdata.worldbank.org/index.php/catalog/dhs#_r=1552916300770&collection=&country=&dtype=&from=1890&page=1&ps=&sid=&sk=india&sort_by=rank&sort_order=desc&to=2018&topic=&view=s&vk=. Accessed March 28, 2018.
World Bank Indicators (2015) Fertility rate, total (births per woman) 2015. https://data.worldbank.org/indicator/SP.DYN.TFRT.IN?locations=IN. Accessed March 28, 2018.
Yang, Juhua (2012) Fertility squeeze and gender bias: a quantitative and qualitative analysis of birth planning policy and sex ratio at birth in China. Working Paper.
Yang, Shuzhang and Guangzhou Wang (2006) Parity-sex progression ratio. *Population Research* 30(2), 10–19.

Cite this article: Aksan A-M (2021). Son preference and the fertility squeeze in India. *Journal of Demographic Economics* **87**, 67–106. https://doi.org/10.1017/dem.2020.5

https://doi.org/10.1017/dem.2020.5 Published online by Cambridge University Press

https://doi.org/10.1017/dem.2020.5 Published online by Cambridge University Press

Appendix A

Table A1. Logit regressions: second-born male

	(1) 1992	(2) 1992	(3) 1998	(4) 1998	(5) 2005	(6) 2005	(7) 2015	(8) 2015
1st born male = 1	0.00994	−0.138	−0.0162	−0.136	−0.0681***	−0.121	−0.0731**	−0.311***
	(0.0186)	(0.0945)	(0.0189)	(0.0899)	(0.0206)	(0.0752)	(0.0225)	(0.0914)
Average PSU children born per woman		−0.0757***		−0.0424†		−0.0130		−0.0412
		(0.0223)		(0.0229)		(0.0284)		(0.0364)
1st born male = 1 × Average PSU children born per woman		0.0470		0.0393		0.0225		0.126**
		(0.0297)		(0.0288)		(0.0319)		(0.0441)
Observations	64,716	64,716	66,593	66,593	65,812	65,812	63,750	63,750
AIC	88,955.0	88,943.4	91,900.3	91,899.6	96,406.2	96,409.2	88,641.3	88,628.0
BIC	89,309.1	89,315.6	92,264.6	92,282.0	96,797.2	96,818.4	89,085.4	89,090.2

Observations are weighted using sampling probability weights. All regressions include state fixed effects and control variables. Clustered standard errors in parentheses.
†$p < 0.10$,
*$p < 0.05$,
**$p < 0.01$,
***$p < 0.001$.

https://doi.org/10.1017/dem.2020.5 Published online by Cambridge University Press

Table A2. Logit regressions: third-born male

	(1) 1992	(2) 1992	(3) 1998	(4) 1998	(5) 2005	(6) 2005	(7) 2015	(8) 2015
1st 2 born male = 1	0.0168	−0.480**	−0.0454	−0.578***	−0.0480	−0.352**	−0.212***	−0.895***
	(0.0310)	(0.161)	(0.0308)	(0.153)	(0.0357)	(0.130)	(0.0417)	(0.171)
1st 2 born—1 male and 1 female = 1	0.0183	−0.305*	−0.0299	−0.282*	−0.0328	−0.237*	−0.149***	−0.591***
	(0.0272)	(0.145)	(0.0275)	(0.140)	(0.0304)	(0.111)	(0.0349)	(0.148)
Average PSU children born per woman		−0.113**		−0.103**		−0.0864*		−0.187***
		(0.0370)		(0.0373)		(0.0422)		(0.0558)
1st 2 born male = 1 × Average PSU children born per woman		0.154**		0.170***		0.124*		0.335***
		(0.0496)		(0.0479)		(0.0529)		(0.0774)
1st 2 born—1 male and 1 female = 1 × Average PSU children born per woman		0.100*		0.0811†		0.0842†		0.220**
		(0.0447)		(0.0438)		(0.0461)		(0.0678)
Observations	48,087	48,087	47,536	47,536	42,744	42,744	37,304	37,304
AIC	66,906.5	66,895.5	66,006.1	65,992.7	64,884.9	64,879.7	47,753.3	47,724.7
BIC	67,257.8	67,273.1	66,365.6	66,378.5	65,266.1	65,286.8	48,179.6	48,176.6

Observations are weighted using sampling probability weights. All regressions include state fixed effects and control variables. Clustered standard errors in parentheses.
†$p < 0.10$,
*$p < 0.05$,
**$p < 0.01$,
***$p < 0.001$.