Shifting social norms around son preference: Impact of a large media campaign in India

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Abstract

Policy measures that seek to address son preference through restrictions on the tools of sex-selective abortions, without addressing the underlying causes, have been found to generate negative welfare consequences for unwanted surviving girls. Unlike these top-down supply-side measures, demand-side measures that focus on increasing the demand for girls can mitigate the adverse welfare consequences that emerge from the birth of unwanted girls. We study the impact of an intervention designed to address son preference, as manifesting in a male-biased sex ratio. The intervention, implemented in India between 2015-18, had both supply-side and demand-side elements, through tighter policing of illegal sex-selective abortions and a mass media campaign designed to increase the perception of the value of a female child. We exploit variation in the timing of exposure to the programme across Indian districts as well as quasi-exogenous variation in the sex of the firstborn child to identify the impact of the programme and find that it led to an increased proportion of female births as well as a reduction in the gender gap in mortality. The main mechanism is an increase in health investments in daughters, such as breastfeeding and vaccinations.

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1 Introduction

A key policy challenge in many deeply patriarchal societies characterised by a strong social norm of son preference is how to tackle the problem of “missing women”, as exemplified by a male-biased sex ratio (Sen 1992). Missing females at birth largely result from the use of foetal sex-screening technologies such as ultrasound following by sex-selective abortions. Many governments around the world, including in China, India, South Korea, Taiwan and Vietnam, have responded to this challenge with top-down measures such as outright bans on abortions, bans on sex-selective abortions or restrictions on the use of ultrasound and other foetal sex-screening technologies. These supply-side measures seek to limit access to the tools of prenatal sex selection without necessarily affecting the demand for prenatal sex-selection: as such, while they may lead to a more balanced sex ratio, in the absence of changes to underlying social norms, they may also displace prenatal discrimination to postnatal margins (Goodkind 1996). Families that cannot practice prenatal sex selection may begin to discriminate against unwanted girls, either directly or indirectly, leading to worsening gender gaps in health and educational outcomes. In contrast, demand-side measures seek to shift the demand for male and female children by changing underlying social norms. Such policies have previously taken the form of mass media campaigns that inform, educate and advocate for more progressive gender norms, and shift perceptions about the economic and social value of daughters. Could such policies mitigate the adverse welfare consequences of increased female births by increasing investments in surviving daughters?

In this paper, we examine the impact of a policy intervention with both supply and demand-side elements on the probability of female births as well as child mortality and health outcomes. The supply-side elements seek to restrict access to sex-selective abortions while the demand-side elements comprise a mass media campaign that promotes the value of daughters and encourages families to invest in their health and education. A policy with both supply-side and demand-side elements would lead to an increase in female births, but the impact on the gender gap in health outcomes would be theoretically ambiguous. If the supply-side measures dominate, we would expect to see an increase in female births with worsening health outcomes for “unwanted” daughters relative to sons. In particular, the increased births of unwanted girls can lead to reduced investments in their human capital and well-being, worsening their health outcomes. This could take place on account of open discrimination against girls, compared to boys. It could also result from families resorting to the use of fertility stopping rules where they keep having children until they achieve a desired number of sons. In this case, girls are disproportionately born into larger families, where they face increased competition for sibling resources, leading a widening gender gap in health outcomes across the entire population.

However, demand-side measures could mitigate these adverse consequences without affecting the births of female children. If demand-side measures dominate, the negative impact on the gender gap in health outcomes could be reversed entirely, especially if families increase health investments in their daughters.

We estimate treatment effects of a mass media campaign launched in India – the Beti Bachao Beti Padhao (Save Girls, Educate Girls) programme – aimed at increasing levels of gender sensitisation, promoting the perception that girls are as valuable as boys, creating incentives for female education and reducing gender discrimination. This campaign was rolled out along with several supply-side interventions strengthening the implementation of the existing legal restrictions on sex screening and sex-selective abortions. The rollout of the campaign was staggered across districts: during the first phase, 100 districts were covered, during the second phase, an additional 160 districts were introduced, and finally the campaign was extended to the entire country in the third phase.

We exploit intertemporal and spatial variation in the rollout of the programme to estimate its impact on the probability of female births, as well as the gender gap in mortality and health outcomes. We additionally use quasi-exogenous variation in the sex of the firstborn child to identify the impact of the intervention on families that are most intensively affected by the intervention. Previous research has found that Indian families have strong elder son preference (Jayachandran, Anukriti, Bhalotra and Tam 2021), and while prenatal sex selection is not common at the first birth order, families with firstborn females are significantly more likely to resort to sex-selective abortions than families with firstborn males. Using this variation, we estimate a triple difference estimator that estimates the impact of the intervention on the gender gap in health outcomes in firstborn female families treated by the programme, compared with firstborn male families treated by the programme.

We find that while the programme did lead to an increase in the proportion of female births, it led to a decrease in the gender gap in mortality in intensively treated families. Moreover, this decrease in mortality for girls was driven by increasing investments in children, such as increased breastfeeding and vaccinations of both pregnant women and their children. Fertility continued to increase in firstborn female families treated by the ban, suggesting that families did resort to the fertility stopping rule to achieve a desired number of sons when they could no longer make use of sex-selective abortions. However, the increased competition of sibling resources did not disadvantage female children in particular, as families were also more likely to invest in their daughters, compared to their sons. Our approach is robust to potential bias emerging from non-random placement of the programme across districts, and we find no evidence of pre-existing trends that could be driving our results.

We provide the first estimates of the treatment effects of the mass media campaign on mortality, health outcomes and health investments, allowing us to comment on the relative efficacy of a supply-side policy – the ban on access to sex-selective abortions – compared to a policy that incorporates demand-side interventions seeking to shift the underlying level of son preference. There is a growing body literature that points to the potential for demand-side interventions to change hardwired social preferences and norms (Anderson et al., 2017). Jensen and Oster (2009) find that the introduction of cable television in rural India increased women's autonomy in making fertility decisions and decreased the acceptability of domestic violence and son preference. In the context of Brazil, La Ferrara et al. (2012) find that access to television, especially soap operas, significantly lowers fertility among women. Using a school-based intervention in the state of Haryana in India that engaged adolescents in classroom discussions about gender equality, Dhar et al. (2018) find that the programme made children's attitudes more supportive of gender equality particularly among boys. Ours is the first analysis of the impact of media campaign on the gender gap in child mortality. While our analysis is limited by the fact that we only observe data on children between 1 and 5 years after the implementation of this programme, our preliminary results are promising and suggest the importance of demand-side elements, such as media-based efforts, to shift social norms through gender sensitisation efforts. This has important insights for policy design for countries struggling to reduce pervasive and deep rooted gender discrimination and to address the problem of missing women.

The rest of the paper is organised as follows. Section 2 presents the background on the media intervention campaign, as well as theoretical motivation on the likely impact of the policy on gender bias; Section 3 describes the data and presents descriptive statistics; Section 4 presents the empirical strategy; Section 5 presents the results along as well as a discussion of potential mechanisms; and Section 6 concludes the discussion.

2 Background to the programme

The context for this study is India, where sex-ratios have long been male-biased. Census data from 2011 put the sex ratio at 943 females per 1000 males, with considerable variation across states from 877 females per 1000 males in Haryana to 1084 females per 1000 males in Kerala. A major cause of the male-biased sex ratio has been the widespread use of ultrasound technology since the 1980s (Bhalotra and Cochrane, 2010), which is used to determine the sex of the foetus, followed by the selective abortion of female foetuses. The national and state governments reacted to the worsening sex ratio by banning sex-selective abortions and placing restrictions on access to ultrasounds through a series of legislations passed between 1989 and 2002.

These bans were found to be effective in increasing female births (Nandi and Deolalikar 2015), but they have also led to worsening gender gaps in human capital outcomes due to relatively reduced investments in girls compared to boys. Lower investments were the outcome of outright discrimination, as in the case of lower educational investments leading to widening gender gaps in educational outcomes (Rastogi and Sharma 2022). They also resulted from increasing family size as families, in the absence of access to abortion, begin to rely on the fertility stopping rule to achieve a desired number of sons (Dasgupta and Sharma 2021). Girls were disproportionately born into larger families after the bans on abortions were enacted, and suffered from increased sibling competition for resources, leading to lower investments in early childhood health, higher mortality and worsened health outcomes.

In 2015, a mass media campaign called the Beti Bachao Beti Padhao (BBBP) programme, or the "Save daughters, educate daughters" programme, was launched in some districts in India with the aim of shifting the social norm of son preference, while simultaneously strengthening the policing of illegal sex-selective abortions. While the supply-side elements of the programme would be anticipated to lead to more female births, the demand-side elements of the programme could mitigate the adverse effects of discrimination against "unwanted" girls, either by directly increasing investments in girls, or, indirectly, by reducing fertility as families are encouraged to be satisfied with the birth of a daughter and not to persist in trying for a desired number of sons.

The programme was designed to address male-biased sex ratios, and to promote women's empowerment and gender equality, by preventing sex-selective abortions, ensuring the survival and protection of the girl child and reducing gender gaps in access to education. The programmes's goals by 2018-19 were to: improve the sex ratio in selected districts by 2 percentage points every year, reduce the gender differentials in under-5 mortality rates from 7 percentage points to 1.5 percentage points, improve female nutrition by reducing the number of anaemic and underweight girls, and increase the enrollment of girls in secondary education to 82 percent. The mass communication campaign involves spreading awareness and disseminating information through radio jingles in Hindi and regional languages, televised messages, community engagement through mobile exhibition vans, social media and field publicity using hand-outs, brochures, text messages on mobile phones in English, Hindi and regional languages. [1]

Ensuring the effective implementation of the ban on sex detection and sex-selective abortions was also a part of the BBBP programme. Local officials were made to monitor the sex ratio at birth and register all births through the Civil Registration System. All pregnancies were to be registered along with the provision of antenatal care (ANC) and postnatal services. All genetic laboratories and clinics conducting any preconception and prenatal diagnostic counselling or tests were to be registered and a complete database of complaints about violations of the ban was to be maintained. Sting operations were conducted to unearth the illegal practice of sex selection (GoI, 2019).

As it focused on both a more stringent implementation of the ban as well as encouraging changes in social norms through advocacy and media campaigns, the programme has both supply-side and demand-side elements. This provides a unique setting to examine the efficacy of legal bans when coupled with demand-side interventions that can change the underlying son preference that drives gender discrimination.

The budgetary allocation of funds from 2015-2018 was in excess of Rs 11 billion (GoI, 2019). In a short period of time the BBBP programme has become very well known: a recent survey of 14 states finds nearly 88 per cent of respondents were aware of the programme (Sinha et al., 2020).

The programme was initially launched in 100 districts in 2015 (Phase 1), and was expanded to 61 additional districts in 2016 (Phase 2). It was expanded to the rest of the country by 2018 (Phase 3). The initiative mainly involved a mass communication campaign targeted at shifting social norms and perceptions about the worth of the girl child, as well as some additional actions in selected districts where the child sex ratio had increased in favour of males between 2001 and 2011. The mass media campaign was launched at the national level, with focused interventions in programme districts. So far there has been very limited research on the impact of the programme. Gupta et al. (2018) examine the short-run impacts of the program in Haryana but they are only able to compare outcomes from before and after the implementation of the programme. They find a significant improvement in the sex ratio at birth in favour of females when analysing data from 2005-2016 for the state of Haryana.

[1] Other actions included a renewed focus on the enforcement of the ban on sex detection and sex-selective abortions. State governments and district-level officials were also asked to improve data collection of birth registrations and the district-level sex ratio at birth through the existing network of health workers and local government structures. Some other measures in the context of health include improvements in the prenatal and postnatal care of mothers, and the provision of counselling to ensure the equitable care of female infants, as well as the training of front-line health workers to make them more sensitive to these concerns. On the educational front, measures include universal enrollment of females in school and construction of toilets specifically for the use of females, as well as the integration of gender-related awareness in the educational curriculum, and gender-sensitisation training of police and judicial personnel.

We use the staggered timing in the roll out of the program across districts to estimate if the relative mortality and health investments for girls improve in districts exposed to the programme, compared to girls in untreated districts. Additionally, we exploit the quasi-exogenous assignment of the gender of the firstborn child to estimate the impact of the programme on the gender gap in relatively intensively treated families – those with firstborn females – compared to less intensively treated families. To our knowledge, these are the first causal estimates of the impact of the programme on child health outcomes.

3 Data

To examine the impacts of the BBBP programme, we pool retrospective birth data from two rounds of the National Family Health Survey(NFHS), a national household survey conducted in 2015-16 and 2019-2020. For fertility outcomes, we use pooled data from retrospective birth histories of all women aged between 15-49 years to construct a dataset of all births that take place in a ten year period between 2011-2020. This dataset includes over 3.2 million mother-year observations on almost 600,000 unique women.

For child health and mortality outcomes, we organise the data at the level of the child. The data on child birth and mortality include 640,00 child observations. In addition, we consider anthropometric outcomes of approximately 390,000 children as well as health investments such as vaccinations of 200,000-300,000 children.

The data also includes a rich set of mother and household characteristics including mother's age, mother's age at childbirth, whether the mother has completed primary education, total children ever born to the mother, religion, caste, whether the household is located in an urban area, household wealth index and total number of members in the household.

4 Empirical Strategy

4.1 Impact on female births

We first examine the impact of the BBBP programme on the proportion of female births. We estimate the following equation:

$$
\begin{aligned}
Y_{bmdy} &= \beta_0 + \beta_1 \text{FirstbornFemale}_m + \beta_2 (\text{Treat} \times \text{FirstbornFemale})_{mdy} \\
&+ \eta \text{X}_{mdy} + \delta_d + \gamma_y + \eta_d Year + \lambda_b + \epsilon_{bmdy} \qquad (1)
\end{aligned}
$$

where Y_{ibmdy} is, for any given year y, the current proportion of female births out of all births for a mother m from district d, and at parity b. $Treat_y$ takes the value one for all years after the programme has been implemented in that district and zero otherwise. *Firstborn Female*$_m$ is defined at the mother level, taking the value one if mother m has a firstborn female child and zero otherwise. We include birth order fixed effects (λ_b) and district-year fixed effects (δ_{dy}). Further, the estimation includes X_m, a vector of socioeconomic and demographic characteristics comprising of mother's age, whether the mother completed primary education, mother's religion, mother's caste, whether the house is located in an urban area, number of members in the household, household wealth index, and sex composition of adults in the household. Standard errors are clustered by district. The main coefficient of interest, β_2, estimates the change in the proportion of female children out of all children for a mother that we observe after exposure to the BBBP programme, in firstborn female families compared to firstborn male families.

The dataset is a full fertility history of 440,000 mothers through the period 2011-2020, including only those mothers who had at least one birth through this period. We drop mothers who had their first child more than 20 years before the start of the survey so as to limit the recall period. The results of this estimation are presented in Table 1. We find that mothers in treated districts were more likely to have female children at birth orders of greater than one, while there is no significant impact in the probability of a female being born as a firstborn child. This effect is driven by firstborn female families: such families are 1.8 percentage points more likely to have female birth than a male birth. Women who were surveyed in treated districts were more likely to have a greater proportion of female children than women who were surveyed in non-treated districts.

In an alternative specification, we use a binary variable for whether child i born to mother m in year y is female. The coefficient of interest on the interaction term between $Treat$ and $FirstbornFemale$ now captures the increased probability of a female birth in treated firstborn female families relative to treated firstborn male families. The results are presented in Table 2. As above, we find that the probability of female births relatively increases after exposure to the programme among firstborn female families, and this difference is significant at the 1% level.

This provides evidence that exposure to treatment had the greatest impact among families that were more likely to resort to favour sons - those which had a firstborn female child. These families would have been more likely to resort to sex-selective abortions in the absence of the programme. However, exposure to the programme reduces their access to illegal ultrasounds and abortions, while also potentially shifting the social norms that drive their preferences for the birth of sons, leading to lower demand for sons. It is in such families that female children are now more likely to be born.

4.2 Impact on the gender gap in mortality

Having established that the programme did lead to a rise in female births, particularly among firstborn female families, we next test whether the increase in the proportion of female children led to a change in the gender gap in mortality and health outcomes. We use the quasi-exogenous variation in the sex of the firstborn child to identify the effect on the gender gap in health outcomes by estimating the following triple difference specification:

$$\begin{aligned} Y_{ibmdty} &= \beta_0 + \beta_1(\text{Treat} \times \text{FirstbornFemale} \times \text{Female})_{ibmhdty} \\ &+ \beta_2(\text{Treat} \times \text{FirstbornFemale})_{mty} + \beta_3(\text{Treat} \times \text{Female})_{ibmdty} \\ &+ \beta_4(\text{FirstbornFemale} \times \text{Female})_{ibmdty} + \beta_5\text{Female}_{ibmdty} \\ &+ \beta_6\text{FirstbornFemale} + \eta\text{X}_{my} + \delta_{dy} + \tau_t + \rho_b + \epsilon_{ibmdty} \end{aligned} \tag{2}$$

where Y_{ibmdty} captures a range of mortality and health outcomes for child i of birth order b born to mother m in district d in month t and year y. $Treat_{dty}$ takes the value one if the child is born after the programme was implemented in the district and zero otherwise. $Female_{ibmdty}$ takes the value one if the child's sex is female and zero otherwise. $Firstborn\ Female_m$ is defined at the mother level, taking the value one if mother m of child i has a firstborn girl child and zero otherwise. We include the triple interaction of these three variables as well as all pairwise interactions between them. We include district-birth year fixed effects (δ_{dy}), birth month fixed effects (τ_t), and birth order fixed effects (ρ_b). Further, the estimation includes X_{my}, a vector of socioeconomic and demographic characteristics comprising mother's age at birth, mother's age at the time of the survey, whether the mother completed primary education, mother's weight for height, mother's religion, mother's caste, whether the house is located in an urban area, number of members in the household, household wealth index and sex composition of adults in the household. Standard errors are clustered by state.

β_1, the coefficient on the interaction of triple interaction term between *Treat*, *Firstborn Female* and *Female*, is our coefficient of interest and captures whether the gender gap in health outcomes of children born into firstborn female families are differentially affected by the programme compared to the gender gap in children born into firstborn male families. Since the $Treat$ variable varies both over time and across district, this coefficient is identified through four sources of variation: spatial and intertemporal variation in exposure to treatment, quasi-exogenous variation in the sex of the firstborn child, and the sex of the child. Districts were selected for the programme based on their pre-programme sex ratios, with those districts with the lowest numbers of females per males being selected first. We control for this variation through the use of district-year fixed effects which capture level differences in the districts assigned to different treatment phases.

We use two measures of child mortality: neonatal mortality (if a child died before completing 1 month) and infant mortality (if the child died before completing 1 year). We are not able to consider under-five mortality (if a child died before completing 5 years) because most children have not been fully exposed to five years of the programme. The sample includes children born between 2011 and 2020 in all states.

In addition to mortality, we also consider health outcomes, including (i) a set of objective biomarkers such as height for age and weight for age; (ii) indicators for health investments that could affect these biomarkers such as ante-natal care (ANC) visits, whether a mother has received tetanus shots while pregnant and breastfeeding duration; and iii) the vaccine status of the children for a range of diseases including measles, hepatitis B, polio, DPT, and BCG. A detailed note on variable definitions and construction can be found in Appendix A.

5 Results

5.1 Impact of the BBBP intervention on child mortality and health outcomes

The results of the impact of the BBBP intervention on mortality are presented in Table 3. The estimated coefficient on the triple interaction term between *Treat*, *Firstborn Female* and *Female* is negative and significantly different from zero at the 5% level in the case of neonatal mortality. The coefficient is also negative and also of a very similar size for infant mortality, though this is not significantly different from zero. This indicates that the gender gap in child mortality declined in families that were most intensively treated by the programme, while controlling for exposure to the programme.

This is a striking result, particularly when compared to the estimated impact of supply-side restrictions on abortion, which find that bans on prenatal sex selection led to an increase in the gender gap in mortality. While the BBBP programme also incorporates supply-side elements, it prominently includes a mass media campaign that was aimed at shifting social norms around the perceived desirability of daughters. As a result, treated families are both more likely to have more daughters and more likely to treat them better as well.

We also examine anthropometric outcomes, such as height-for-age (HFA), weight-for-age (WFA) and body mass index (BMI) that relate to health in Table 4. Unlike with mortality, there is no significant reduction in the gender gap in these health outcomes. This could be because these observations are on children who are between the ages of 0-5 years, which means the older children have not yet been fully exposed to the treatment. Another way to interpret these results is that despite an increase in female births caused by the programme, there is no negative effect on the observed health outcomes of female children. On the other hand, a ban on sex-selective abortions – a pure supply-side measure to tackle son preference – caused a rise in the gender gap in height and weight outcomes (Dasgupta and Sharma 2022).

We next examine the gender gap in investments in children in Tables 5 and 6. We consider some prenatal investments such as antenatal visits and tetanus shots, as well as several postnatal investments such as months of breastfeeding and whether a child received a routine vaccination. We find that mothers of girls are 6.7 percentage points more likely to receive a tetanus shot than mothers of boys, when comparing the gender gap among treated firstborn female and treated firstborn male families. This difference is statistically significant at the 1% level. Similarly, the gender gap in months of breastfeeding also narrows by 1.48 months. Among common vaccinations received by children, the coefficients on the triple interaction term are positive and significant for DPT vaccines, and positive and statistically significantly different from 0 for others, like measles, hepatitis B, polio and BCG.

In sum, in stark contrast to results that find a worsening of female child mortality health outcomes as a result of supply-side measures such as bans on prenatal sex-selection, a policy with demand-side elements is able to mitigate and outright reverse some of these adverse consequences. We observe lower mortality outcomes for girls, potentially driven by increased parental investments such as in breastfeeding and in vaccinations. These indicate significant benefits from the implementation of a gender-equity focused policy with a strong demand-side component.

5.2 Pre-intervention trends

We test for the possibility of pre-intervention trends that could be biasing our results by restricting our analysis to children born before their districts were exposed to the programme. We interact the variables $FirstbornFemale$ and $Female$ with an indicator for the each of five years preceding treatment. The results of this estimation are in Table 7. All the coefficients on the lagged terms are insignificantly different from 0.

5.3 Impact on fertility

To investigate the effects of BBBP programme on fertility, we use a similar estimation framework as equation (1). Specifically, we test if fertility increases relatively more in firstborn female families as compared to firstborn male families after the implementation of the policy. We run the following estimation:

$$\begin{aligned} Y_{mdy} &= \beta_0 + \beta_1 \text{Treat}_{mdy} + \beta_2 \text{FirstbornFemale}_{mdy} + \\ &+ \beta_3 (\text{Treat} \times \text{FirstbornFemale})_{mdy} + \beta_4 X_{mdy} + \delta_{dy} + \epsilon_{mdy} \end{aligned} \tag{3}$$

where Y_{mdy} is an indicator for fertility for mother m in district d in year y. We include district-birth year fixed effects (δ_{sy}) and a set of controls X_{msy} that includes mothers education, mothers age, total children born, religion, caste, living in an urban area, total number of household members, and family structure. The main coefficient of interest is β_3, the coefficient on the interaction between an indicator for a firstborn female family and the treatment indicator.

The results are in Table 8. We find that fertility continued to increase in firstborn female families even after exposure to treatment. In other words, the programme leads to an increase in the likelihood of female births, as well as an increase in the probability of any birth. This suggests use of the fertility stopping rule by families who continue to have children to increase their chances of achieving a desired level of sons. However, unlike in the case of pure supply-side measures, we do not observe a worsening in health outcomes of girls relative to boys. Presumably, any negative effects from the increased competition for sibling resources that are disproportionately faced by girls are cancelled out or even reversed by the benefits from increased care and investments in female child health.

6 Discussion

We examine the impact of a large-scale intervention to tackle son preference that focuses both on supply-side measures that reduce access to sex-selective abortions, as well as a mass media intervention that seeks to shift social norms and increase the demand for girls. We find that the policy does lead to an increase in female births, it also leads to an improvement in female health outcomes, relative to males. In doing so, the policy reverses the adverse consequences of having a large number of "unwanted" female births, or the rise in fertility that results from families using the fertility stopping rule to achieve a desired number of sons. This emphasises the importance of incorporating demand-side elements to policies that seek to eliminate gender discrimination, rather than simply focus on top-down approaches that address some tools of discrimination without addressing the underlying causes.

References

Anderson, S., D. Stromberg, and J. Walfogel (Eds.) (2017). *Economic and Social Impacts of the Media*. Elsevier.

Bhalotra, S. R. and T. Cochrane (2010). Where have all the young girls gone? Identification of sex selection in India. IZA Discussion Papers 5381, IZA Institute of Labor Economics.

Dhar, D., T. Jain, and S. Jayachandran (2018). Reshaping adolescents' gender attitudes: Evidence from a school-based experiment in india. Technical report, National Bureau of Economic Research.

GoI (2019). Beti bachao, beti padhao scheme. Implementation guidelines, Ministry of Women and Child Development, Government of India.

Gupta, R., R. Nimesh, G. L. Singal, P. Bhalla, and S. Prinja (2018). Effectiveness of india's national programme to save the girl child: experience of beti bachao beti padao (b3p) programme from haryana state. *Health policy and planning 33*(7), 870–876.

Jensen, R. and E. Oster (2009). The power of tv: Cable television and women's status in india. *The Quarterly Journal of Economics 124*(3), 1057–1094.

La Ferrara, E., A. Chong, and S. Duryea (2012). Soap operas and fertility: Evidence from brazil. *American Economic Journal: Applied Economics 4*(4), 1–31.

Sinha, A., R. K. Jaiswal, R. Sundar, D. P. Singh, A. Alawadhi, V. Rangarajan, and P. Dhawan (2020). An Evaluation of India's Beti Bachao Beti Padhao Scheme. Report 20200802, National Council of Applied Economic Research.

7 Tables

8 Appendix

Table 1: Impact of BBBP on sex ratio

Sex Ratio	Proportion of Female Children						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
(8)	(9)						
	All Birth Orders	All Birth Orders	Birth Order=1	Birth Order>=2	Birth Order>=2	Last Birth Year	Last Birth Year
tab_sexratio_s6_m1_11	tab_sexratio_s6_m1_12						
Treat	-0.00482***	0.00000	0.00431	0.00000	-0.00183	0.00776**	0.00000
0.01447***	0.00000						
	(0.00126)	(.)	(0.00386)	(.)	(0.00860)	(0.00379)	(.)
(0.00553)	(.)						
Firstborn Female		0.62386***		0.63360***			0.39751***
	0.55253***						
		(0.00477)		(0.00413)			(0.00267)
	(0.00505)						
Treat X Firstborn Female		-0.05997***		0.01806***			0.01478***
	0.07908***						
		(0.00248)		(0.00283)			(0.00274)
	(0.00373)						
Observations	2521361	2521361	648081	648076	245388	402693	402682
422716	422711						
Mean of Dep. Variable	0.4946	0.4946	0.5064	0.5064	0.4826	0.5216	0.5216
0.4853	0.4853						
SD	0.387	0.387	0.398	0.398	0.500	0.315	0.315
0.377	0.377						

Standard errors in parentheses
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Proportion of female children is the ratio of the total number of female children to the total number of children born to a mother. Columns 1 and 2 include all births during 2011-2020. Column 3 includes all first order births. Columns 4 and 5 include all higher-order births during 2011-2020. Columns 6 and 7 estimates the *sex ratio* of children ever born to a mother as measured in the year of her last birth. *Treat* takes the value 1 if BBBP has been implemented in a mother's district in a particular year, 0 otherwise. *Firstborn Female* takes the value 1 if firstborn child is female, 0 otherwise. All estimations control for mother's age in a particular year, whether the mother has completed primary education, religion, caste, whether the house is located in an urban area, number of members in the household, household wealth index and sex composition of adults in the household. Additionally columns 2, 5 and 7 include district fixed effects, year fixed effects, district-year fixed effects and parity fixed effects while columns 1, 3, 4 and 6 include district fixed effects, year fixed effects and parity fixed effects. Columns 6 and 7 does not include parity fixed effects instead controls for the total number of children born. The sample is restricted to woman who gave birth to at least one child between 2011 and 2020. Standard errors are clustered at the district level.

Table 2: Impact of BBBP on probablity of child being female

	(1)	(2)
	Female	Female
Treat	0.00510	0.00000
	(0.00520)	(.)
Firstborn Female		0.37238***
		(0.00476)
Treat X Firstborn Female		0.01646***
		(0.00445)
Observations	648081	648076
Mean of Dep. Variable	0.4783	0.4783
SD	0.500	0.500

Standard errors in parentheses
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Female takes the value 1 if a child is female, 0 if boy. *Treat* takes the value 1 if BBBP has been implemented in a mother's district in a particular year, 0 otherwise. *Firstborn Female* takes the value 1 if firstborn child is female, 0 otherwise. All estimations control for mother's age in a particular year, whether the mother has completed primary education, religion, caste, whether the house is located in an urban area, number of members in the household, household wealth index and sex composition of adults in the household and additionally includes district fixed effects, year fixed effects, district-year fixed effects and parity fixed effects. Column 1 does not include district-year fixed effects. The sample is restricted to woman who gave birth to at least one child between 2011 and 2020. Standard errors are clustered at the district level.

Table 3: Impact of BBBP on gender gap in child mortality by firstborn female family

	(1) NNM	(2) IMR
Treat	-0.00154	-0.00245
	(0.00348)	(0.00481)
Female	-0.00821***	-0.00955***
	(0.00093)	(0.00123)
Treat x Female	0.00511**	0.00540
	(0.00208)	(0.00335)
Firstborn Female	-0.00468***	-0.00826***
	(0.00101)	(0.00120)
Treat x Firstborn Female	0.00272	0.00044
	(0.00196)	(0.00290)
Female x Firstborn Female	0.00527***	0.00921***
	(0.00131)	(0.00174)
Treat x Female x Firstborn Female	-0.00691**	-0.00662
	(0.00280)	(0.00430)
Observations	666030	586235
Mean of Dep. Variable	0.0272	0.0420
SD	0.163	0.201

Standard errors in parentheses
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Note: Neonatal Mortality (*NNM*) takes the value 1 if a child died before completing 1 month, 0 otherwise. Infant Mortality (*IMR*) takes the value 1 if a child died before completing 1 year, 0 otherwise. Children who have not attained these respective ages are excluded from the regression. *Treat* takes the value 1 if BBBP has been implemented in a mother's district in a particular year, 0 otherwise. *Female* takes the value 1 if the child's sex is female, 0 otherwise. *Firstborn Female* takes the value 1 if firstborn child is female, 0 otherwise. All estimations control for mother's age, mother's age at birth, whether the mother has completed primary education, religion, caste, whether the house is located in an urban area, number of members in the household, household wealth index and sex composition of adults in the household and additionally includes birth order fixed effects, district fixed effects, birth year fixed effects, birth month fixed effects, and district-year fixed effects. The sample is restricted to children born between 2011 and 2020. Standard errors are clustered at the district level.

Table 4: Impact of BBBP on gender gap in child anthropometric outcomes by firstborn female family

	(1) HFA	(2) WFA	(3) BMI
Treat	0.07194	-0.00749	-0.07261*
	(0.04644)	(0.02804)	(0.03750)
Female	0.09009***	0.04129***	-0.00987
	(0.01354)	(0.00990)	(0.01299)
Treat x Female	0.09078***	0.12159***	0.03471
	(0.02912)	(0.01894)	(0.02485)
Firstborn Female	0.05501***	0.05455***	0.03339***
	(0.01233)	(0.00938)	(0.01031)
Treat x Firstborn Female	0.02754	0.02607	0.01507
	(0.02571)	(0.01740)	(0.02173)
Female x Firstborn Female	-0.06357***	-0.05207***	-0.01636
	(0.01919)	(0.01337)	(0.01772)
Treat x Female x Firstborn Female	-0.05020	-0.03601	-0.01790
	(0.04214)	(0.02502)	(0.03087)
Observations	393747	397843	390261
Mean of Dep. Variable	-1.4054	-1.5095	-0.8170
SD	1.748	1.263	1.469

Standard errors in parentheses
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Note: *HFA* refers to the child's height for age z-score. *WFA* refers to the child's weight for age z-score. *BMI* is the BMI of the child. *Treat* takes the value 1 if BBBP has been implemented in a mother's district in a particular year, 0 otherwise. *Female* takes the value 1 if the child's sex is female, 0 otherwise. *Firstborn Female* takes the value 1 if firstborn child is female, 0 otherwise. All estimations control for child's age, mother's age, mother's age at birth, whether the mother has completed primary education, religion, caste, whether the house is located in an urban area, number of members in the household, household wealth index and sex composition of adults in the household and additionally includes birth order fixed effects, district fixed effects, birth year fixed effects, birth month fixed effects, and district-year fixed effects. The sample is restricted to children born between 2011 and 2020. Standard errors are clustered at the district level.

Table 5: Impact of BBBP on gender gap in child health investments by firstborn female family

	(1) ANC	(2) Duration Breastfed	(3) TET
Treat	-0.15714	-1.22676***	0.02420
	(0.12873)	(0.22100)	(0.01469)
Female	0.11194***	0.10819	0.01915***
	(0.04223)	(0.11888)	(0.00715)
Treat x Female	-0.02473	-0.33986**	-0.04215***
	(0.08514)	(0.14679)	(0.01097)
Firstborn Female	-0.10938***	0.69577***	0.01967***
	(0.03696)	(0.09141)	(0.00634)
Treat x Firstborn Female	0.04126	-0.89831***	-0.02876***
	(0.08871)	(0.11408)	(0.00978)
Female x Firstborn Female	0.78235***	-1.48203***	-0.03482***
	(0.06197)	(0.15507)	(0.01066)
Treat x Female x Firstborn Female	-0.16652	1.48115***	0.06729***
	(0.15520)	(0.18065)	(0.01579)
Observations	666030	353605	336271
Mean of Dep. Variable	12.2653	15.2376	1.9256
SD	8.131	12.514	0.743

Standard errors in parentheses
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Note: *ANC* refers to the number of antenatal care visits that took place in utero, top-coded at 20. *TET* refers to the number of tetanus injections the mother received before birth. *Breastfeeding Duration* refers to the duration of breastfeeding in months. *Treat* takes the value 1 if BBBP has been implemented in a mother's district in a particular year, 0 otherwise. *Female* takes the value 1 if the child's sex is female, 0 otherwise. *Firstborn Female* takes the value 1 if firstborn child is female, 0 otherwise. All estimations control for mother's age, mother's age at birth, whether the mother has completed primary education, religion, caste, whether the house is located in an urban area, number of members in the household, household wealth index and sex composition of adults in the household and additionally includes birth order fixed effects, district fixed effects, birth year fixed effects, birth month fixed effects, and district-year fixed effects. The sample is restricted to children born between 2011 and 2020. Standard errors are clustered at the district level.

Table 6: Impact of BBBP on gender gap in child vaccination outcomes by firstborn female family

	(1) Measles	(2) Hepatitis B	(3) Polio	(4) DPT	(5) BCG
Treat	0.00367	0.02106**	0.00749	0.02842***	-0.00310
	(0.01129)	(0.00870)	(0.00691)	(0.00777)	(0.00613)
Female	0.00205	0.00299	-0.00017	-0.00010	-0.00231
	(0.00374)	(0.00494)	(0.00449)	(0.00410)	(0.00327)
Treat x Female	-0.00273	0.00611	0.00008	-0.00206	0.00208
	(0.00679)	(0.00848)	(0.00967)	(0.00795)	(0.00689)
Firstborn Female	0.01332***	0.01734***	0.00966**	0.01301***	0.00830***
	(0.00348)	(0.00438)	(0.00412)	(0.00317)	(0.00292)
Treat x Firstborn Female	-0.01154**	-0.01030	-0.00015	-0.01495***	-0.00502
	(0.00581)	(0.00706)	(0.00562)	(0.00528)	(0.00378)
Female x Firstborn Female	-0.01507***	-0.01664**	-0.00963	-0.01128**	-0.00647
	(0.00526)	(0.00693)	(0.00635)	(0.00574)	(0.00493)
Treat x Female x Firstborn Female	0.01316	0.00170	0.00130	0.01835**	0.00419
	(0.00868)	(0.01175)	(0.01075)	(0.00933)	(0.00881)
Observations	342101	200071	218025	295201	344147
Mean of Dep. Variable	0.6422	0.7525	0.8780	0.8430	0.9162
SD	0.479	0.432	0.327	0.364	0.277

Standard errors in parentheses
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Note: *Measles* takes the value 1 if the child had received vaccine against measles, 0 otherwise. *Hepatitis B* takes the value 1 if the child had received all the four shots of Hepatitis B vaccine, 0 otherwise. *Polio* takes the value 1 if the child had received all the four shots of oral Polio vaccine, 0 otherwise. *DPT* takes the value 1 if the child had received all the three shots of DPT vaccine, 0 otherwise. *BCG* takes the value 1 if the child had received BCG vaccine, 0 otherwise. *Treat* takes the value 1 if BBBP has been implemented in a mother's district in a particular year, 0 otherwise. *Female* takes the value 1 if the child's sex is female, 0 otherwise. *Firstborn Female* takes the value 1 if firstborn child is female, 0 otherwise. All estimations control for child's age, mother's age, mother's age at birth, whether the mother has completed primary education, religion, caste, whether the house is located in an urban area, number of members in the household, household wealth index and sex composition of adults in the household and additionally includes birth order fixed effects, district fixed effects, birth year fixed effects, birth month fixed effects, and district-year fixed effects. The sample is restricted to children born between 2011 and 2020. Standard errors are clustered at the district level.

Table 7: Impact of BBBP on gender gap in child mortality by firstborn female family

	(1) NNM	(2) IMR
2012 x Female x Firstborn Female	-0.00695*	-0.00651
	(0.00372)	(0.00465)
2013 x Female x Firstborn Female	0.00302	0.00368
	(0.00380)	(0.00460)
2014 x Female x Firstborn Female	0.00252	0.00473
	(0.00400)	(0.00493)
2015 x Female x Firstborn Female	-0.00406	-0.00621
	(0.00454)	(0.00641)
2016 x Female x Firstborn Female	-0.00570	-0.00230
	(0.00520)	(0.00628)
2017 x Female x Firstborn Female	-0.00353	-0.00403
	(0.00508)	(0.00635)
Observations	534773	497923
Mean of Dep. Variable	0.0274	0.0402
SD	0.163	0.196

Standard errors in parentheses
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Note: Neonatal Mortality (*NNM*) takes the value 1 if a child died before completing 1 month, 0 otherwise. Infant Mortality (*IMR*) takes the value 1 if a child died before completing 1 year, 0 otherwise. Children who have not attained these respective ages are excluded from the regression. *Treat* takes the value 1 if BBBP has been implemented in a mother's district in a particular year, 0 otherwise. *Female* takes the value 1 if the child's sex is female, 0 otherwise. *Firstborn Female* takes the value 1 if firstborn child is female, 0 otherwise. All estimations control for mother's age, mother's age at birth, whether the mother has completed primary education, religion, caste, whether the house is located in an urban area, number of members in the household, household wealth index and sex composition of adults in the household and additionally includes birth order fixed effects, district fixed effects, birth year fixed effects, birth month fixed effects, and district-year fixed effects. The sample includes only hcildren born in untreated districts between 2011 and 2017. Standard errors are clustered at the district level.

Table 8: Impact of BBBP on fertility

	(1) Any Birth	(2) Any Birth
Treat	0.00641***	0.00000
	(0.00238)	(.)
Firstborn Female		0.01822***
		(0.00076)
Treat X Firstborn Female		0.00440***
		(0.00121)
Observations	3331250	3331250
Mean of Dep. Variable	0.1940	0.1940
SD	0.395	0.395

Standard errors in parentheses
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Any Birth takes the value 1 if a child is born to the mother in a particular year, 0 otherwise. *Treat* takes the value 1 if BBBP has been implemented in a mother's district in a particular year, 0 otherwise. *Firstborn Female* takes the value 1 if firstborn child is female, 0 otherwise. All estimations control for mother's age in a particular year, whether the mother has completed primary education, religion, caste, whether the house is located in an urban area, number of members in the household, household wealth index and sex composition of adults in the household and additionally includes district fixed effects, year fixed effects, district-year fixed effects and parity fixed effects. Column 1 does not include district-year fixed effects. The sample is restricted to woman who gave birth to at least one child between 2011 and 2020. Standard errors are clustered at the district level.

Appendix A Variable definitions

1. Indicators for health investments

   (a) The number of antenatal visits the women had while the child was in utero. The value of these visits were topcoded at 20+ visits, while the the children whose mothers did not go for antenatal care were coded as 0. According to WHO recommendations, there should be a minimum of eight antenatal visits to decrease perinatal mortality and improve women's experience of care.

   (b) This variable reports if and how many tetanus toxicoid vaccinations were given to mother while the child was in utero for children born in three to five years before the survey. According to WHO recommendations, in case the mother is not previously vaccinated or in the case of unknown vaccination status of mother, she should be given two doses of tetanus toxicoid vaccination one month apart, with the second dose given at least two weeks before the delivery.

   (c) Breastfeeding refers to months of breastfeeding for the children born in three to five years before the survey including the cases where (a) the child's mother was still breastfeeding at the interview time and (b) the child had been breastfed until his/her death. On a population basis, exclusive breastfeeding for 6 months is the best way of feeding infants, and after that infants should be continued with breastfeeding for up to 2 years of age or beyond along with complementary foods.

2. Set of objective biomarkers

   (a) Height for age z-score captures the height for age z-score value for surviving children born in three to five years before the survey. According to the WHO global database on child growth and malnutrition a height for age z score between -2 & -3 is characterized as moderate chronic malnutrition, while that below -3 corresponds to severe chronic malnutrition.

   (b) Weight for age z-score captures the weight for age z-score value for surviving children born in three to five years before the survey. Low child weight for age indicates acute/chronic malnutrition. According to WHO global database on child growth and malnutrition weight for age z-score between -2 & -3 corresponds to moderate malnutrition, while that below -3 corresponds to severe malnutrition.