1 Introduction

In this paper we analyse the effect of the Integrated Child Development Services (ICDS) programme initiated by the Indian government in 1975 on the gender difference in child immunisations. The ICDS programme aims to provide child immunisation services, supplementary feeding and nutrition, and childcare education to young mothers across India via childcare centres, or *anganwadi* centres. ICDS is currently the largest early childcare programme in the world, with over 748,000 centres operating across the country as of 2006 according to the Ministry of Women and Child Development (MWCD).[1] We examine whether the programme reduces or intensifies the bias against female children with regard to immunisation. We also examine how the effect of the increase in healthcare supply on gender bias varies over time. The policy implications of this research are crucial given the twin objectives of increasing access to healthcare and reducing inequality in developing countries. Increasing access to healthcare for children is an important part of the Millennium Development Goals. But if the effect of increased access is to worsen and make permanent already existing inequalities, then additional solutions need to be found to prevent this from occurring.

This paper makes two contributions to the literature. The first is an examination of the effect of a permanent increase in healthcare supply on the level of gender discrimination in immunisation over time. The response of the gender difference in immunisation to supply shocks has been addressed previously in Oster (2009a). Oster analyses the effect of a continuous increase in healthcare supply on mean differences in immunisation between boys and girls at the village-level, using number of health-camp visits to the village as a measure of healthcare supply. The results show a non-monotonic response with an initial increase in gender inequality that eventually dissipates. The ICDS programme provides a similar downward shock to the cost of immunisations for households. However, Oster (2009a) exploits continuous variation in supply at a given point in time. We, on the other hand, are interested in how variation in the length of time households are exposed to a supply shock of fixed intensity affects their response. We find that exposure to the programme widens the male-female disparity in vaccinations by 0.022-0.025 immunisations. The increase in gender difference is also greatest for households that have been treated for the least amount of time; a result that can be considered analogous to Oster's finding higher gender differences in households that receive fewer health-camp visits.

[1] Figures and information on the ICDS programme are provided by the Ministry of Women and Child Development (MWCD), Government of India at `http://wcd.nic.in/`

The second contribution of this paper is to the literature regarding the ICDS programme itself. To the best of our knowledge, this is the first paper that attempts to evaluate the effects of the ICDS programme on gender discrimination in immunisation levels. The results have significance for policy, given the large size of the programme and its consequent potential to alter gender allocations in child health investments.

Section 2 presents background information on the causes of gender bias against women in India, and details on the implementation of the ICDS programme. Section 3 describes the data and the empirical methodology used in our analysis. The results from the analysis and robustness checks are presented in Section 4. Finally we discuss the implications of our results for policy and the gender ratio in India in Section 5.

2 Background

Discrimination against Indian female children in the provision of nutrition and health care is often identified in the economics literature as the root cause of the large gender ratio imbalance observed in the Indian population. In the 2001 Census of India, the female-male gender ratio (the number of females divided by the number of males in the population) was as low as 0.861 and 0.874 in the states of Haryana and Punjab respectively, and an even lower 0.821 in the national capital Delhi (Census of India (2001)). The literature on intra-household resource allocation in India has established a pattern of discrimination against female children in the provision of nutrients (for e.g. see Rose (1999); Behrman (1988); Sen and Sengupta (1983); Rosenzweig and Schultz (1982)). The demography literature has also shown evidence that the bias against girls in India affects household fertility patterns, as well as provision of health investments such as breastfeeding and immunisation (Mishra et al. (2004); Pande (2003); Marcoux (2002); Clark (2000); Arnold et al. (1998); Das (1987)). This discrimination manifests in excess female child mortality, which creates the observed imbalance in the population gender ratio. [2]

Gender bias in India is driven by a combination of sociocultural and economic factors which determines the net utility from each gender. Aside from the consumption value of children, parents broadly derive economic utility from their

[2] Anderson and Ray (2008) provide an alternative point of view to the literature, arguing that the gender ratio in India is biased across *all* age groups and not just in the child population. However, they argue that the bulk of the ”missing women” in India can be explained by excess female deaths due to infectious disease; an assertion which they also admit does not rule out discrimination against girls in childhood which increases susceptibility to disease at older ages.

children from wage earnings, labour in agricultural production, and security in illness and old age (Bardhan (1974); Arnold et al. (1998)). Sons traditionally stay with the family and work the household land or contribute to household earnings, and bring additional labour when the wife joins the household. There is an additional premium brought to the household by male children in the form of a very large dowry at the time of marriage. In contrast to sons, daughters bring several costs to the household. The most important of these costs is a very large dowry that is customarily paid by the daughter's family to the husband's family at the time of marriage (Kishor (1995)). This dowry is traditionally very large (up to several multiples of annual earnings), and places a tremendous financial burden on the parents of the female child. Also as mentioned previously, daughters become a part of their husband's household after marriage. Their labour on agricultural land is therefore lost to their natal household. There is an additional large opportunity cost of the time required to arrange a daughter's marriage, and to protect her chastity while she is unmarried.

2.1 ICDS Programme

The ICDS programme was begun in 1975, and is currently the world's largest early child development programme. The programme offers a package of health and educational services to children under the age of 6 and their mothers, all of which are provided at *anganwadi* centres which operate at the level of the village. The services offered include health check-ups, treatment of worms and diarrhea, growth monitoring, provision of micronutrient supplements, health and nutrition education to mothers, preschool education, and child immunisation. Since its inception, the programme has been expanded extensively; the number of administrative blocks covered rose from 33 in 1975, to more than 6,118 in 2006.[3] By 2006, the programme services were being provided to 46.7 million children and 9.5 million pregnant and lactating women through 7.48 million *anganwadi* centres across the country (MWCD (2009)).

The ICDS programme has been evaluated previously with respect to its impact on child nutritional outcomes. Deolalikar (2004) finds a 5% reduction in the probability of boys being underweight in villages with an ICDS centre, but finds no significant association for girls. Lokshin et al. (2005) however, find no significant effect of the programme on nutritional outcomes upon using propensity score matching methods. The impact of the programme on child immunisation has yet to be empirically investigated.

[3] see MWCD reference above.

An important aspect of the programme that is relevant for our analysis is how it is placed regionally. Specifically, since we are analysing the effects of the programme on gender bias in immunisation, we must ensure that the programme is not targeted to areas with systematically different levels of gender bias in immunisation levels. According to the official programme policy, the targeted beneficiaries are children up to the age of 6, pregnant women, and mothers who belong to the poorest sections of society and reside in backward rural and tribal areas or urban slums. While there is no explicit evidence that the programme also targets populations where female children are discriminated against with respect to immunisation or other health inputs, it is possible that it does so incidentally if such gender discrimination is more severe among poorer communities.

To further alleviate concerns regarding bias from programme targeting, we run placebo regressions and implement a difference-in-differences procedure to validate our results. We also perform checks on the probability of programme coverage at the state level. The results of these analyses are presented in our robustness checks section.

3 Data and Empirical Strategy

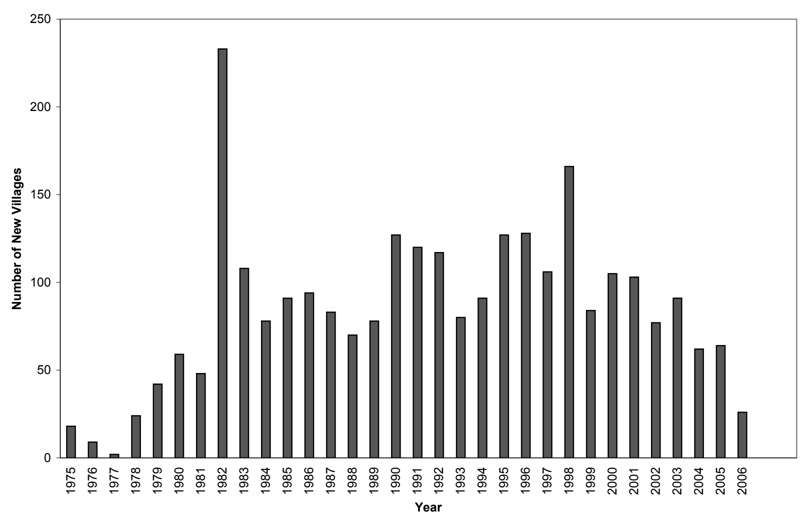
We use the 2005-06 round of the Measure DHS survey in India, which contains detailed birth histories, child health, and maternal health information for over 100,000 women aged 15-49 years. Detailed immunisation information is available for all children born during 2001-2006. Information on whether the household is covered by an ICDS centre and the number of years the ICDS centre has been open is also collected as part of the survey. To preserve homogeneity of the sample we only consider women who are currently married and have only been married once. We also only retain women who have had a total of six births or less as the sample size shrinks rapidly by birth order after this point.[4]

From the sample described above, Figure **??** shows the number of new villages introduced to the ICDS programme every year since the inception of the programme in 1975. In the first six years of the programme except for 1980, less than 50 new villages were covered by the programme annually. 1982 appears to be unusual year, with 233 villages in our sample being introduced to programme coverage. The number of villages included in the programme annually then becomes more stable. The set of vaccines supplied by the government and delivered

[4] The full sample for which vaccination data is available consists of 51,555 children and 36,850 women. We remove 735 women who are not currently married, and an additional 762 women who have been married more than once. Restricting the sample of children born during 2001-2006 by birth order leaves us with 33,685 women (or households) and 46,879 children.

to households via the ICDS programme was only finalised in 1985. The Expanded Programme on Immunisation (EPI) begun in 1978, introduced a supply of the BCG, DPT, and oral Polio vaccines to urban areas. The Measles vaccine was added in 1985 with the initiation of the Universal Immunisation Programme (UIP), which aimed to provide a supply of vaccines to all parts of the country (WHO (2005)). As all the nine vaccines currently provided under the ICDS programme have only been part of the delivery package since 1985, we remove women who reside in villages that were treated before 1986 from the sample, in order to preserve comparability in the treatment package received across households. This leaves us with a twenty year period to analyse programme effects. This length of time period for our analysis also serves to create symmetric categories by length of coverage, and remove the early outlier years of programme implementation from the sample.

Figure 1: New Villages Included in ICDS Programme Per Year (1975-2006)



Notes: Figures presented for sample trimmed by marriage and birth order

3.1 Immunisations and Household Characteristics

There are nine total vaccinations on which data is collected for children born during the period 2001-2006, the six years leading up to the survey. These are the BCG vaccine, Measles vaccine, four rounds of the Polio vaccine, and three rounds of the DPT vaccine. We use the total number of vaccinations received by children born during these years as our measure of child health investment. Table **??** presents mean vaccination levels by child gender for each of the six birth cohorts in the sample trimmed by birth order and marriage. For every cohort barring that born in 2006, the mean number of vaccinations received by boys is greater than the mean for girls, with boys receiving 0.10-0.30 more vaccinations on average. For the cohorts born in 2001, 2003, and 2005, the difference in means is statistically significant. The lack of a similar difference in immunisation levels for the youngest cohort is not surprising, as vaccinations are acquired over time.

Table 1: Mean Difference Tests of Vaccinations By Gender

Birth Year	Boys	Girls	Difference	SE	Fraction of Girls	Frequency
2001	6.29	6.08	0.21**	0.076	0.473	6,575
2002	6.28	6.18	0.10	0.064	0.484	9,414
2003	6.37	6.24	0.13*	0.062	0.471	9,578
2004	6.51	6.39	0.12	0.062	0.478	9,294
2005	5.57	5.27	0.30**	0.062	0.483	9,656
2006	2.54	2.52	0.02	0.102	0.495	2,362

Notes: Means are reported for the sample trimmed by marriage and birth order. The nine vaccinations covered are the BCG vaccine, Measles vaccine, four rounds of the Polio vaccine, and three rounds of the Diphtheria (DPT) vaccine. ** Significant at 1% ; * Significant at 5%.

Table **??** reports the total number of villages and households in the trimmed sample, along with figures on ICDS coverage. There are a total of 3,056 villages, of which 1,999 have an ICDS centre. 27,495 households and 38,435 children make up the final sample. Table **??** shows descriptive statistics of important determinants of parental demand for child immunisation from the trimmed sample. Mother's educational attainment takes integer values from 0 to 5, where 0 indicates no education and 5 indicates education beyond secondary school. Birth order is measured in ascending order, with smaller numbers denoting older births. Other variables include the succeeding and preceding birth intervals in months, mother's age in years at the time of the child's birth, and the mother's report of the child's smallness at the time of birth. These variables help us to ascertain the child's health endowment, which is a crucial determinant of health investment. Finally, we also consider various sibling composition variables such as previous living sons and previous living children, which have been shown to affect demand for child health investments in India given the prevailing bias against female children.

In Table **??**, we present the differences in means and proportions of various household and mother characteristics by ICDS coverage from the estimation sample. Nearly all the differences are highly significant, illustrating that the programme

Table 2: India DHS 2005-06 - Coverage of ICDS Programme

	Villages	Households	Children
Covered	1,999	18,574	26,251
Not Covered	1,057	8,921	12,184
Total	3,056	27,495	38,435
ICDS 0-5 Years	423	4,151	5,916
ICDS 6-10 Years	589	5,624	8,027
ICDS 11-15 Years	535	4,876	6,871
ICDS 16-20 Years	452	3,923	5,437
Total	1,999	18,574	26,251

Notes: Figures are reported for the trimmed sample used for regression analysis. Years of coverage are reported conditional on being covered by an ICDS centre.

Table 3: India DHS 2005-06 - Descriptive Statistics

	Mean	S.D.	Minimum	Maximum
Mother's Educational Attainment	1.80	–	0	5
Mother's Current Age	26.14	4.91	15	49
Previous Living Sons	0.56	0.81	0	5
Previous Living Children	1.21	1.24	0	5
Birth Order	2.35	–	1	6
Mother's Age at Birth	23.41	4.80	11	47
Smallness at Birth	2.99	–	1	5
Preceding Birth Interval	27.40	13.06	9	159
Succeeding Birth Interval	32.52	18.86	9	263

Notes: Statistics are reported for the trimmed sample used for regression analysis. Preceding and succeeding birth intervals are reported conditional on there being a preceding or succeeding birth respectively.

is targeted to low income households. Mothers in ICDS-covered villages are much more likely to live in a rural area, to have married younger, to have given birth earlier, and to live in a house with a dirt or clay floor. They are also much less likely to have completed primary or secondary education, or live in a household that owns a radio, motorcycle, car, or refrigerator. We therefore control for several household wealth indicators in our analysis. While we cannot be certain that we have included all possible wealth indicators that determine selection into treatment, we present robustness checks to reduce concerns of bias from any omitted wealth factors.

Table 4: Tests of Difference in Means and Proportions by ICDS Coverage

	Non-ICDS	ICDS	Difference	S.E.
Mother's Age at Marriage	18.72	17.69	1.03**	0.040
Years of Marriage	7.43	7.78	-0.35**	0.049
Mother's Age at First Birth	24.49	23.88	0.61**	0.062
Child Birth Weight (Grams)	2835.18	2838.34	-3.16	11.32
Rural	0.256	0.768	-0.512**	0.005
Mother Completed Primary Education	0.613	0.485	0.128**	0.005
Mother Completed Secondary Education	0.226	0.092	0.134**	0.004
Muslim	0.189	0.156	0.033**	0.004
Mud/Clay/Earth Floor	0.181	0.376	-0.195**	0.005
Cement Floor	0.388	0.262	0.126**	0.005
Household Owns Refrigerator	0.316	0.128	0.188**	0.004
Household Owns Motorcycle	0.300	0.159	0.141**	0.004
Household Owns Radio	0.354	0.313	0.041**	0.005
Household Owns Television	0.649	0.425	0.224**	0.005
Household Owns Bicycle	0.444	0.501	-0.057**	0.005
Household Owns Car	0.075	0.025	0.050**	0.002

Notes: Means and proportions of mother and household characteristics are reported for the trimmed sample used for regression analysis. ** Significant at 1% ; * Significant at 5%

3.2 Empirical Strategy

To estimate the first order effect of the programme on gender bias in immunisation, we use the following specification:

$$I_{ih} = \beta_1 girl_{ih} + \beta_2 ICDS_h{}^* girl_{ih} + \beta_3 ICDS_h + \gamma_1 X_{ih} + \eta_h + \theta_i + \varepsilon_{ih} \quad (1)$$

I_{ih} is the number of vaccinations received by child i in household h, where each mother in the sample is treated as a single household. The dummy variable $girl_{ih}$ equals 1 when child i is a girl. $ICDS_h$ is a dummy variable for whether the household is covered by an ICDS centre. In addition to these regressors, we include a vector of child-specific and household covariates X_{ih} and a household (or mother) fixed effect η_h. Child-specific unobservables are included in θ_i, and ε_{ih} is an idiosyncratic error term. The beta coefficients in (**??**) will allow us to ascertain the effect of child gender on child immunisation level, and the marginal effect of

ICDS policy coverage on any prevailing gender bias in immunisation. The X_{ih} vector contains various other household and child-specific variables that may influence the number of vaccinations child i receives, such as number of previous children and previous sons, a dummy variable for whether there is at least one previous living son, birth order, and mother's education and age. In addition to controlling for birth order and sibling gender composition, we also include a dummy variable for whether the child in question is the first-born child to the household. This is to control for the documented parental bias in favour of the firstborn child.

To capture the effect of child gender on demand for immunisation without any bias or confounding effects, we must attempt to control for the child's health endowment as best as we can. To do this we include the variables discussed previously, namely preceding and succeeding birth intervals, mother's age at birth, and mother's report of child smallness. While other studies have relied on birth weight rather than reported birth size, we argue that the latter is the correct variable to use as it is the parents' perception of the child's health endowment that influences their decision to immunise. Finally, to control for the effects of household wealth on immunisation and the probability of ICDS centre coverage, we include several household wealth indicators such as whether the household owns a refrigerator, a radio, a bicycle, or a car, the materials from which the house is constructed, the form of toilet in the house, and whether the household has electricity. We also include the DHS household standard of living index which is calculated using several indicators of household wealth.[5] Regional effects are controlled for by including state and village fixed effects.

It is possible that there are still omitted household unobservables that influence both gender bias and the level of child immunisation. Including the household (or mother) fixed effect η_h allows us to control for these potential sources of bias, but doing so also mitigates the household response to the price shock that is driven by other unobservables such as traditional attitudes or personal biases. As this is precisely the household response we want to capture, including the household fixed effect may dampen our coefficient of interest. We therefore present results from estimations both with and without household fixed effects. There is also a concern that child-specific unobservables in θ_i bias the coefficients of interest. These unobservables are part of the combined error term $u_{ih} = \theta_i + \varepsilon_{ih}$ which is possibly correlated with the regressors. We minimise this possible source of bias by including child-specific regressors that control for the health endowment at birth and birth-year dummy variables. It is in any case unlikely that child-specific heterogeneity is

[5]The index is calculated by aggregating scores assigned to several household assets such as livestock, agricultural land, and consumer durables. For more details see the National Family and Health Survey 3 (NFHS3) supplemental documentation.

correlated with the probability of ICDS coverage once we control for regional and household variables. It is also reasonable to assume that child-specific heterogeneity is uncorrelated with child gender, once we control for health endowments.

3.3 Dynamic Effect of Supply Shock on Gender Inequality

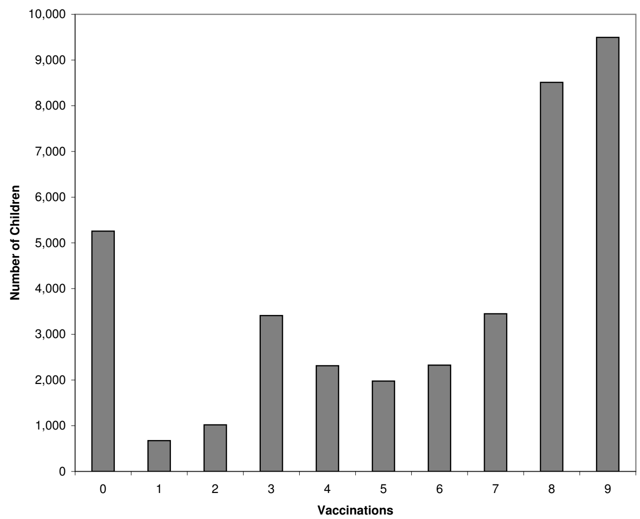
Logically, the response of the gender difference in immunisation to a permanent price shock should vary with duration of exposure to the shock. Lags in information dissemination or lack of awareness may prevent households from immediately taking advantage of the new supply of child vaccines. Or it is possible that when the new supply becomes available, the initial response of households is much larger than in later periods due to pent-up demand. To examine whether the programme effect on the gender difference is sensitive to duration of exposure, we expand the specification in (**??**) to include the second order effects of time as follows:

$$
\begin{aligned}
I_{ih} \quad = \quad & \beta_1 girl_{ih} + \beta_2 ICDS_h{}^* girl_{ih} \\
& + \beta_3 ICDS_h{}^* ICDS_{h,6-10}{}^* girl_{ih} \\
& + \beta_4 ICDS_h{}^* ICDS_{h,11-15}{}^* girl_{ih} \\
& + \beta_5 ICDS_h{}^* ICDS_{h,16-20}{}^* girl_{ih} \\
& + \gamma_1 X_{ih} + \eta_h + \theta_i + \varepsilon_{ih}
\end{aligned} \qquad (2)
$$

Here, $ICDS_{h,6-10}$ is a dummy variable which equals 1 if the ICDS centre has been open for 6-10 years. Similarly, $ICDS_{h,11-15}$ and $ICDS_{h,16-20}$ are dummy variables which equal 1 when the ICDS centre has been open 11-15 years or 16-20 years respectively. The omitted category is ICDS centres that have been open for 0-5 years. We also estimate the specification dividing length of treatment into four-year and two-year categories, again using the children most recently treated as the omitted category. Each dummy variable for duration of coverage is also included separately without interactions in the specification with state fixed effects.

We estimate the specifications in (**??**) and (**??**) using a Poisson event count data model. While this procedure accounts for the discrete nature of the outcome variable, the frequency of children by number of immunisations received, shown in Figure **??**, does not approximate the Normal distribution. There is also overdispersion in the data. We correct for this using a quasi-maximum likelihood procedure which allows for the computation of robust standard errors, and produces consistent estimates relying only on the conditional mean assumption even if the Poisson distribution is misspecified (Wooldridge (1999)). This procedure conveniently deals with the overdispersion problem, as no variance function is required to be specified. We also implemented OLS regressions to verify the results from

Figure 2: Number of Children by Vaccinations Received



Notes: Figures presented for the trimmed regression sample

the Poisson regressions. The main findings do not change between the procedures, so we discuss only the magnitudes of the results from the Poisson regressions.[6] Results are reported for each specification with state, village, and household fixed effects. The state fixed effects procedure is useful to obtain an estimate of the main programme effect, which gets absorbed in the fixed effects in the latter two procedures. As the ICDS centre coverage for households is reported at the level of the village (the primary sampling unit) within each state, we cluster the standard errors at the village level. The results are discussed in the next section.

4 Results

Tables **??**, **??**, and **??** present results from the estimation of (**??**) and (**??**) with state fixed effects, village fixed effects, and household fixed effects respectively.[7] The

[6]The OLS results are available from the author upon request. Negative Binomial regressions were also attempted to correct for the over-dispersion, but convergence proved difficult. Allison and Waterman (2002) have also shown the procedure not to be a ”true” fixed effects model in the majority of cases.

[7]With regard to the household fixed effect specification, the coefficients are however not readily comparable to those in the village and state fixed effect specifications as the estimation procedure only uses the sub-sample of households with more than two births in the years 2001-2006. These

first column in each table reports the coefficient on the gender dummy. In all three tables the coefficient is negative and significant, indicating that on average girls receive 0.02 fewer vaccinations than boys. The ICDS programme effect, also reported in the first column of the state fixed effects table, is to increase vaccinations by an average of 0.046 per child. While this estimate appears small given the possible nine vaccinations each child can receive, it should be viewed in light of evidence from India and other countries that each incremental immunisation significantly reduces the probability of child mortality (Lee (2005); Bawah et al. (2009)). Hence, a positive and significant programme effect can be interpreted as having an impact on child survival. The treatment and gender interaction term in the third column is negative in all three specifications, and significant in the state and village fixed effect specifications. The results suggest that the supply shock from the ICDS programme leads to boys receiving 0.022-0.026 more vaccinations on average than girls. This is 48-58% of the above estimate of the programme effect.

Columns (3)-(5) in Tables **??**, **??**, and **??** report the results after disaggregating the treatment effect by duration of coverage in years using three-way interaction terms. Column (3) divides the duration of treatment into five-year categories, column (4) into four-year categories, and column (5) into two-year categories. In each column the programme-gender interaction is the residual category, representing the shortest period of treatment. This is 0-5 years of treatment in column (3), 0-4 years of treatment in column (4), and 0-2 years of treatment in column (5). In Table **??** we find a strongly significant, negative coefficient on the programme-gender interaction in columns (3), (4), and (5). This indicates that girls receive significantly fewer vaccinations than boys when the ICDS centre has been present in the village for a short period of time. In each column we also find that girls receive significantly more vaccinations than boys when they reside in villages that have been covered by the programme for more than 12-16 years. This is evidence of boys being vaccinated first in the initial years of coverage, and girls being vaccinated later. However, this specification does not control for regional effects at the level of the village, which could be correlated with programme coverage, so we move to the results in Table **??**.

The results from the village fixed effects specification in Table **??** also show that girls receive significantly fewer vaccinations than boys in the initial years of programme coverage in columns (3) and (4). The size of the significant coefficients

households are likely to be qualitatively different from those that are excluded. We were also unable to achieve convergence during the household fixed effects estimation when including mother's age at birth, most likely due to lack of variation or collinearity with other regressors. However we estimated the specification again including this variable but without robust standard errors to simplify the algorithm, and this did not change the values of the other coefficient estimates. Hence we can be reasonably confident that they are not biased by the exclusion of this regressor.

on the programme-gender interaction is actually larger in these columns compared to those in Table **??** once village-level regional effects that influence the programme impact are controlled for. The same is true for the interaction term for 16-20 years

Table 5: Marginal Effects of ICDS Coverage: State Fixed Effects

Number of Vaccinations									
	(1)		(2)		(3)		(4)		(5)
Girl	-0.021**	Girl	-0.005	Girl	-0.005	Girl	-0.005	Girl	-0.005
	(0.007)		(0.008)		(0.008)		(0.008)		(0.008)
ICDS	0.046**	ICDS	0.057**	ICDS	0.057**	ICDS	0.057**	ICDS	0.058**
	(0.011)		(0.013)		(0.013)		(0.013)		(0.013)
		ICDS*Girl	-0.025*	ICDS*Girl	-0.040**	ICDS*Girl	-0.048**	ICDS*Girl	-0.046*
			(0.011)		(0.015)		(0.015)		(0.022)
				ICDS*ICDS 6-10*Girl	0.016	ICDS*ICDS 5-8*Girl	0.021	ICDS*ICDS 3-4*Girl	-0.003
					(0.014)		(0.019)		(0.028)
				ICDS*ICDS 11-15*Girl	0.018	ICDS*ICDS 9-12*Girl	0.021	ICDS*ICDS 5-6*Girl	0.023
					(0.015)		(0.019)		(0.023)
				ICDS*ICDS 16-20*Girl	0.029*	ICDS*ICDS 13-16*Girl	0.041*	ICDS*ICDS 7-8*Girl	0.016
					(0.013)		(0.019)		(0.026)
						ICDS*ICDS 17-20*Girl	0.034	ICDS*ICDS 9-10*Girl	0.031
							(0.019)		(0.027)
								ICDS*ICDS 11-12*Girl	0.005
									(0.026)
								ICDS*ICDS 13-14*Girl	0.022
									(0.035)
								ICDS*ICDS 15-16*Girl	0.052*
									(0.023)
								ICDS*ICDS 17-18*Girl	0.018
									(0.029)
								ICDS*ICDS 19-20*Girl	0.044
									(0.027)
Observations	32,468		32,468		32,468		32,468		32,468
States	29		29		29		29		29

Notes: Robust standard errors are reported in parentheses. Additional regressors include mother's educational attainment, mother's current age, education and age interaction terms, birth order and sibling composition variables, religion and caste fixed effects, birth year fixed effects, and child health endowment variables. ** Significant at 1% ; * Significant at 5%.

Table 6: Marginal Effects of ICDS Coverage: Village Fixed Effects

				Number of Vaccinations					
	(1)		(2)		(3)		(4)		(5)
Girl	-0.021**	Girl	-0.004	Girl	-0.004	Girl	-0.004	Girl	-0.005
	(0.005)		(0.008)		(0.008)		(0.008)		(0.008)
		ICDS*Girl	-0.026*	ICDS*Girl	-0.046**	ICDS*Girl	-0.055**	ICDS*Girl	-0.037
			(0.011)		(0.016)		(0.018)		(0.021)
				ICDS*ICDS	0.023	ICDS*ICDS	0.033	ICDS*ICDS	-0.039
				6-10*Girl	(0.017)	5-8*Girl	(0.020)	3-4*Girl	(0.032)
				ICDS*ICDS	0.016	ICDS*ICDS	0.023	ICDS*ICDS	0.025
				11-15*Girl	(0.019)	9-12*Girl	(0.021)	5-6*Girl	(0.027)
				ICDS*ICDS	0.042*	ICDS*ICDS	0.039	ICDS*ICDS	0.006
				16-20*Girl	(0.019)	13-16*Girl	(0.021)	7-8*Girl	(0.025)
						ICDS*ICDS	0.052*	ICDS*ICDS	0.013
						17-20*Girl	(0.022)	9-10*Girl	(0.027)
								ICDS*ICDS	-0.006
								11-12*Girl	(0.029)
								ICDS*ICDS	-0.009
								13-14*Girl	(0.030)
								ICDS*ICDS	0.040
								15-16*Girl	(0.026)
								ICDS*ICDS	0.029
								17-18*Girl	(0.030)
								ICDS*ICDS	0.037
								19-20*Girl	(0.029)
Observations	32,380		32,380		32,380		32,380		32,380
Villages	2,943		2,943		2,943		2,943		2,943

Notes: Robust standard errors clustered at the village level are reported in parentheses. Additional regressors include mother's educational attainment, mother's current age, education and age interaction terms, birth order and sibling composition variables, religion and caste fixed effects, birth year fixed effects, and child health endowment variables. ** Significant at 1% ; * Significant at 5%.

Table 7: Marginal Effects of ICDS Coverage: Household Fixed Effects

	Number of Vaccinations								
	(1)		(2)		(3)		(4)		(5)
Girl	-0.020*	Girl	-0.005	Girl	-0.004	Girl	-0.005	Girl	-0.004
	(0.010)		(0.015)		(0.015)		(0.015)		(0.015)
		ICDS*Girl	-0.022	ICDS*Girl	-0.068**	ICDS*Girl	-0.064*	ICDS*Girl	-0.069*
			(0.016)		(0.023)		(0.026)		(0.032)
				ICDS*ICDS	0.066**	ICDS*ICDS	0.045	ICDS*ICDS	0.012
				6-10*Girl	(0.026)	5-8*Girl	(0.030)	3-4*Girl	(0.046)
				ICDS*ICDS	0.040	ICDS*ICDS	0.046	ICDS*ICDS	0.019
				11-15*Girl	(0.027)	9-12*Girl	(0.029)	5-6*Girl	(0.042)
				ICDS*ICDS	0.067*	ICDS*ICDS	0.039	ICDS*ICDS	0.076*
				16-20*Girl	(0.027)	13-16*Girl	(0.030)	7-8*Girl	(0.039)
						ICDS*ICDS	0.074*	ICDS*ICDS	0.064
						17-20*Girl	(0.033)	9-10*Girl	(0.039)
								ICDS*ICDS	0.035
								11-12*Girl	(0.040)
								ICDS*ICDS	0.027
								13-14*Girl	(0.043)
								ICDS*ICDS	0.057
								15-16*Girl	(0.038)
								ICDS*ICDS	0.074
								17-18*Girl	(0.047)
								ICDS*ICDS	0.084
								19-20*Girl	(0.043)
Observations	18,358		18,358		18,358		18,358		18,358
Households	8,608		8,608		8,608		8,608		8,608

Notes: Robust standard errors clustered at the village level are reported in parentheses. Additional regressors include mother's educational attainment, mother's current age, education and age interaction terms, birth order and sibling composition variables, religion and caste fixed effects, birth year fixed effects, and child health endowment variables. Mother's age at birth is omitted from the regressors. ** Significant at 1% ; * Significant at 5%.

of coverage in column (3). In columns (3) and (4) it appears that girls have levels of immunisation similar to boys in villages that have been covered by the programme for approximately 5-15 years, and receive more vaccinations than boys in villages that have been covered for longer. The evidence again points to boys being immunised before girls in the initial years of programme coverage, with girls catching up later. The two-year categories are too small to pick up gender differentials in treatment with village fixed effects, so all the interaction terms are insignificant in column (5).

While we cannot compare the magnitudes of the coefficients from the household fixed effect specification in Table **??** to those in Tables **??** and **??**, within this sub-sample of households we find the same pattern of results. The programme-gender interaction is negative and significant in columns (3), (4), and (5) indicating disproportionately greater immunisation of boys in the initial years of ICDS programme coverage. In villages that have received longer periods of treatment, girls receive either similar or greater levels of immunisation.[8]

4.1 Robustness Checks

Our results are subject to weakness if the ICDS programme is placed first in areas where gender bias in child health investments is most severe. This would mean our results are being driven by reverse causality. There is nothing in the programme literature to indicate that the timing of the introduction of the programme to villages is influenced by regional conditions, including gender bias in health investments. However, to ensure that this is not the case, we run placebo fixed effect OLS regressions with the same specification, but with months of breastfeeding as the dependent variable. There is a well documented gender bias against girls in breastfeeding in India. Implementing these regressions can alleviate the concern that the policy is targeted, and also eliminate the possibility that the results are driven by unobservables that are simultaneously correlated with gender bias and the ICDS coverage variables. The results are presented in Table **??**.[9] The disadvantage to female chil-

[8]It is worth noting that while more boys may receive immunisations than girls due to the programme, enough children of both genders might might be vaccinated such that the *ratio* of boys to girls that are close to full immunisation actually declines. We investigate this by running OLS estimations of the same specifications using the log of vaccinations as the dependent variable rather than the count of vaccinations. The results do not change, indicating that boys also receive more vaccinations as a *percentage* of their previous immunisation level. The results are available from the author upon request

[9]We do not report the results for the specification with duration of treatment divided into two-year categories for the sake of conciseness. None of the programme related regressors are significant in this specification either.

dren in breastfeeding is evident in the results from the specification without the ICDS programme indicators in the first column. However, none of the programme-related regressors are significant in the regressions with months of breastfeeding as the dependent variable. These results support the hypothesis that the programme is not targeted to areas with higher gender bias.

Regarding the concern of omitted wealth variables, there has not been a significant amount of empirical research on the effect of poverty on gender discrimination in India. Within the literature that exists, there is a tentative hypothesis that gender bias is less severe in poorer households (Krishnaji (1987); Miller (1993); Dasgupta (1993)). Others have argued that poverty is not a significant determinant of gender bias (Das Gupta (1987)). Rose (1999) presents evidence that income shocks have a greater adverse effect on female child survival probabilities in rural landless households than in those with land holdings. However this provides little insight on how existing levels of gender bias vary by household wealth status prior to these income shocks. Murthi et al. (1995) find that female child survival probabilities improve with greater household poverty, lending further support to the notion that gender bias is more intense at higher levels of income. However it is difficult to conclusively accept this hypothesis without more empirical investigation.

To alleviate concerns that omitted wealth variables or other unobservables potentially correlated with treatment biasing our own results, we exploit the age differences between the children in our sample to perform a difference-in-differences (DID) robustness check. Since the children are meant to receive all nine vaccinations by the age of nine months, it is highly likely that the majority of older children do not receive immunisations via the ICDS programme in villages where coverage has only recently been introduced. We therefore take a sub-sample of children and divide them into two groups: those who were born in 2004-06 and those born in 2001-02. We then also divide these children into two groups based on their treatment status: those who do not have ICDS coverage and those who received programme coverage beginning in 2004-06. We then take the difference in mean vaccinations across age groups and again across treatment groups to arrive at a DID estimate of the ICDS programme effect. This estimate is based on the assumption that most of the children born in 2001-02 do not receive treatment if they live in villages where the programme was introduced in 2004 or later, making them a suitable control group for comparison with those who are young enough to be treated. We perform the DID procedure separately for boys and girls, and report the results in Table **??**. We find a positive and significant programme effect of 0.46 vaccinations for boys in Panel A, but no effect for girls in Panel B. This is further evidence suggesting that boys are immunised first in the initial years of programme coverage, widening the gender difference in immunisation levels. To control further for within-household differences in education, wealth, and other factors we run the

Table 8: ICDS Coverage and Gender Bias in Breastfeeding

	Months of Breastfeeding			
	(1)	(2)	(3)	(4)
Girl	-0.533** (0.110)	-0.548** (0.189)	-0.548** (0.189)	-0.548** (0.189)
ICDS*Girl		0.022 (0.232)	0.163 (0.342)	0.015 (0.371)
ICDS*ICDS 6-10*Girl			-0.443 (0.370)	
ICDS*ICDS 11-15*Girl			-0.095 (0.389)	
ICDS*ICDS 16-20*Girl			0.099 (0.414)	
ICDS*ICDS 5-8*Girl				-0.047 (0.417)
ICDS*ICDS 9-12*Girl				-0.422 (0.423)
ICDS*ICDS 13-16*Girl				0.361 (0.441)
ICDS*ICDS 17-20*Girl				0.291 (0.473)
Village Fixed Effects	Yes	Yes	Yes	Yes
Observations	32,190	32,190	32,190	32,190
Number of Villages	3,003	3,003	3,003	3,003
R-Squared	0.31	0.31	0.31	0.31

Notes: Robust standard errors clustered at the village level are reported in parentheses. Additional regressors include mother's educational attainment, mother's current age, education and age interaction terms, birth order and sibling composition variables, religion and caste fixed effects, birth year fixed effects, and child health endowment variables. ** Significant at 1% ; * Significant at 5%.

full specification in (**??**) with state fixed effects on both age-group samples and report the results in Table **??**. We still find a gender difference of 0.06 vaccinations in favour of boys due to the programme in the younger sub-sample which is significant at the 10% level after adding the full set of regressors. There is no such difference as a result of programme coverage in the older group. [10]

5 Discussion

Our results indicate that households that receive ICDS coverage have little or no access to child immunisations before the supply shock. The increase in supply and accompanying fall in price therefore cause not only the total number of vaccinations to increase, but also the gender difference in vaccinations. However the gender differential that is created by the increase in supply disappears over time. From the evidence the likeliest mechanism by which the gender disparity increases is that households immunise their sons before their daughters in the initial stages of exposure to the ICDS programme. Daughters and sons are immunised on a more equal basis in areas where the programme has been in operation for longer. Deolalikar (2004) also found similar results for the ICDS programme effect on malnutrition, with boys having 5% less probability of being malnourished but no such effect for girls.

The policy implications of these results are important with regard to commitment and length of planned health interventions. In countries such as India where gender bias in child health investment is endemic, a sudden increase in healthcare supply where availability was previously low may have adverse effects on equality in the short run. However if the increase in supply is sustained over time, equality in health investment levels is reestablished, ostensibly at a higher

[10] We run additional robustness checks, the results of which are not reported. To control for the gender-specific effects of other policies on immunisation, we estimate the specification again after including state indicators interacted with the gender dummy. This is the best we can do without more specific village-level information on additional government policies. After including the state-gender interaction terms the pattern and significance of the results do not change, which allows greater confidence in their validity. To also rule out possible problems due to programme targeting we use data from the previous two waves of the DHS surveys in India to implement an OLS estimation of the number of ICDS centres in each state in 1998-99 on the mean immunisation level, and also the difference in mean vaccinations between girls and boys in each state in 1992-93. To capture the fact that the programme is targeted to low income households, we include the state average as well as the average gender difference in child weight-age percentile in 1992-93. The estimates reveal that ICDS centres appear to be targeted to areas with the lowest mean vaccination and weight-age percentile levels. However neither the gender difference in mean immunisation nor the gender difference in weight-age percentile are significant in determining the number of ICDS centres. The results tables from these additional robustness checks are available from the author upon request.

Table 9: ICDS Coverage and Gender: Vaccinations DID

Panel A: Boys			
	Born 2001-02	Born 2004-06	Difference
No ICDS	6.368** (0.065)	5.655** (0.061)	0.713** (0.089)
ICDS	5.389** (0.161)	5.139** (0.137)	0.250 (0.211)
Difference	0.979** (0.170)	0.516** (0.152)	0.463* (0.229)
Panel B: Girls			
	Born 2001-02	Born 2004-06	Difference
No ICDS	6.240** (0.070)	5.593** (0.063)	0.646** (0.094)
ICDS	5.299** (0.158)	4.586** (0.139)	0.712** (0.209)
Difference	0.940** (0.172)	1.007** (0.156)	-0.066 (0.232)

Notes: Estimations use children in ICDS villages who are only exposed to programme coverage in 2004 or later. Standard errors are in parentheses. ** Significant at 1% ; * Significant at 5%

level of investment than previous to the supply shock. Hence policy interventions such as the ICDS programme which are permanent or sufficiently long lasting can achieve higher child health investment levels with only short term consequences for inequality.

The findings of our analysis mirror those presented in Oster (2009a). Oster examines the effect of a continuous increase in healthcare supply measured by health camps on gender inequality in immunisation in India, and finds that at initial levels of the supply increase inequality worsens. As supply continues to increase the inequality eventually dissipates. In our case, a permanent increase in availability of vaccines via the ICDS programme represents a long term, invariant shock in supply. However the household response to the supply shock over time is identical to the response found by Oster to an increase in supply itself. Both sets of results suggest that households respond to initial and uncertain increases in healthcare supply by insuring their sons against illness first, as boys are seen to be more valuable than girls. Daughters are then later vaccinated when supply increases further or when the initial increase is seen to be more than just temporary.

Table 10: Vaccinations by Age-Group

	Number of Vaccinations	
	Born 2004-06	*Born 2001-02*
ICDS*Girl	-0.062* (0.037)	-0.008 (0.031)
Girl	-0.005 (0.015)	-0.001 (0.011)
ICDS	0.017 (0.033)	-0.020 (0.018)
Observations	5,509	4,519
States	29	29

Notes: Estimations use children in ICDS villages who are only exposed to programme coverage in 2004 or later. State fixed effects are included in the specification. Robust standard errors are in parentheses. *** Significant at 1% ; ** Significant at 5%; * Significant at 10%

The implications of gender biased immunisation rates for the future gender ratio in India have been discussed to some extent in the existing literature. Oster (2009b) argues that gender differences in immunisation are responsible for upto between 20 to 30 percent of the childhood mortality difference between boys and girls, including when each of seven immunisations is analysed separately. Indeed, it is possible that gender discrimination in childhood health inputs are responsible for the missing Indian women in older age groups as calculated in Anderson and Ray (2008). The authors argue that the bulk of these older missing women can be attributed to a gender differential in susceptibility to infectious disease. It is very possible that gender discrimination in childhood immunisations plays a large part in creating this greater vulnerability to disease in adult Indian women. Lee (2005) has shown that immunisation has a tremendous negative impact on child and infant mortality rates in India. The study estimates that after controlling for various selection mechanisms in the demand for health care, full immunisation reduces the probability of child mortality by 80-95% and partial immunisation by 75-80%.[11] Given that our estimated ICDS programme effect is approximately 0.05 vaccinations and that the programme has reached 46 million children, the less optimistic

[11] These estimates may appear large but Bawah et al. (2009) find similar individual impacts of the same immunisations in northern Ghana. They also find that children below age five are more than 50% less likely to die if partially immunised, and 70% less likely if fully immunised.

results from Lee (2005) would mean that the programme has saved 1.72-1.84 million children. [12] Our estimates of gender bias in immunisation increases due to the programme indicate that approximately 70% of these children are boys.

There are of course positive externalities from immunisation in the form of restricted spread of disease which perhaps make the net effect of biased immunisation coverage on mortality less unequal. Nevertheless, given the severity of the gender ratio imbalance in India, it is vital that healthcare policy interventions are designed to deal with such gender discrimination. The results from our analysis suggest that a long term, stable supply of immunisations is an effective means of tackling childhood mortality without worsening the gender ratio further. Temporary or sporadic increases in supply however create a potential trade off between these two concerns, as lower childhood mortality may only be achieved at the cost of a more adverse gender ratio in the population.

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[12] The lower bound assumes that all treated children are partially immunised, and that as a result their survival probability improves by 75%. The upper bound assumes that all treated children are fully immunised, and that their survival probability improves by 80%.

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