Fertility and Child Mortality: Unintended Consequences of Family Formation in India

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November 17, 2008

Abstract

This paper examines how fertility decisions can lead to excess female mortality. In several developing countries, parents tend to follow son-preferring fertility stopping rules. Stopping rules create a distribution of households: smaller households have a high proportion of sons and larger households have a high proportion of daughters. I examine both the household size effect and the sex composition effect with two complementary economic models. The first model shows that even if parents care equally about the survival of their sons and daughters, larger economic benefits of sons relative to daughters cause parents to follow stopping rules. The model predicts that girls are better off with brothers and boys are better off with sisters. Focusing on India, I use the sex outcome of the first pregnancy as a "natural experiment". I find that stopping rules can explain about a quarter of the gap between boy and girl mortality in children of birth order two and higher.

I also examine the Indian institutions of dowry and inheritance rights that favor sons as the possible underlying economic causes of anti-girl bias. Dowries are a large cost for daughters, but are not necessarily a benefit for sons. The second model, in which parents seek to maximize their bequest per son, highlights one testable change in the costs and benefits of daughters: legal reforms that give inheritance rights to women. These reforms cause parents to lower their health investments in both sons and daughters. I use state reforms granting inheritance rights in a difference-in-difference analysis to provide evidence that giving women inheritance rights has the unintended consequence of raising child mortality rates.

Keywords: Fertility; Child Mortality; Dowry; Inheritance; India.

JEL Classification Numbers: I1, I30, J13, J16, O12, O53.

*E-mail: Daniel.Rosenblum@yale.edu. I thank Christopher Udry, Dean Karlan, Mark Rosenzweig, and Paul Schultz for guidance. I thank Prabhat Jha for advice and for providing access to the Indian Reproductive and Child Health Survey II. I also thank Stephan Klasen and Greg Fischer for their valuable comments and suggestions. This paper also benefited from seminar participants at Yale University, as well as participants at a seminar at the Centre for Global Health Research.

1 Introduction

Child mortality in the least developed countries (LDCs) is high. The under five mortality rate in LDCs is over 15% compared to 0.6% in industrialized countries.[1] Furthermore, in most of the countries of South Asia and in China, girls have much higher mortality rates than boys (Fuse and Crenshaw, 2006). For example, Arnold et al. (1998) found that for children aged one to four in India, girls had mortality rates 43% higher than boys. Higher female mortality is of particular concern because males are biologically weaker than females. Thus, without any discrimination against girls, we would expect higher mortality rates amongst boys than girls, as is seen in all developed countries. There are several papers on South Asia that find discrimination against girls in the provision of health resources, yet the reasons for this discrimination are not well understood.

This paper posits two complementary theories with empirical support to explain excess female mortality in countries with a strong son preference. One stresses the importance of sons' larger economic benefits to their parents compared to daughters. The second focuses on the specific costs and benefits of bequeathing land to children. Both models show that fertility decisions, driven by economic incentives, increase the child mortality gap. I am able to test this hypothesis using the sex outcome of a first-birth as a natural experiment and find that stopping rules can account for a large portion of excess female mortality. The inheritance model provides a testable hypothesis on what happens when the costs and benefits of children change. When women are granted inheritance rights, parents who prefer to give their land to their sons will reduce both son and daughter health spending in order to decrease family size and increase the amount of land each surviving son can inherit. I test this hypothesis with a difference-in-difference analysis using state variation in the granting of inheritance rights to women and find that an unintended consequence of providing inheritance rights to women is higher child mortality.

Parents with a strong preference for sons use two methods to affect their fertility: sex-selective abortion and the choice to continue having children until one has a desired number of boys, commonly referred to as a stopping rule. The pervasiveness of son-preferring stopping rules in India has been well-documented (Clark, 2000; Arnold et al., 2002). A son-preferring stopping rule creates a

[1] http://www.unicef.org/statistics/index.html.

distribution of households.[2] Households with few children will have a high proportion of boys, while larger households will have a high proportion of girls. Thus, stopping rules cause girls on average to be in households with more siblings than boys. We always expect girls to end up in households with a higher proportion of girls compared to boys, regardless of the fertility rule. However, given a "large" household of N children, those N children will have a higher proportion of girls under a stopping rule population than under a population of households that have N children of random sex. Hence, a stopping rule increases the expected proportion of girls in a girl's family. The theories presented in this paper are new economic explanations for why parents follow son-preferring fertility stopping rules. Sex-selective abortion has become more frequent in India (Jha et al., 2006), and its use should reduce the number of children in a household and increase the proportion of boys. In a way, sex selective abortion counteracts the effects of stopping rules. However, unlike stopping rules, selective abortion directly reduces the number of women in the population.

Having more siblings than boys, *the household size effect*, causes girls to live in households with fewer resources per child. One new contribution of the paper is deriving the second consequence of stopping rules, *the sex composition effect*, from an economic model of fertility decisions and the future costs and benefits of girls and boys. The higher the proportion of girls in a household, the larger the incentive for parents to discriminate against girls and favor boys. Papers on family formation tend to treat sex composition as exogenous to child outcomes (Jensen, 2003; Angrist and Evans, 1998; Subramanian and Deaton, 1991; Das Gupta, 1987; Muhuri and Preston, 1991; Pande, 2003). Jensen in particular is similar to my own research and focuses on the effect of son-preferring stopping rules on education outcomes in India. Jensen investigates the household size effect but ignores the sex composition effect. As I show, both the number and sex composition of children are endogenous with respect to child outcomes. I contribute to the literature on family formation by taking into account both these effects in my empirical investigations.

The difference in economic costs and benefits of boys versus girls influence decisions on whether to have additional children. These fertility decisions will in turn affect how parents invest in their children. Children are costly. Parents have to feed and clothe them, as well as reduce their labor

[2] See Keyfitz (1968) p.379-384 for a brief exposition on the mathematics of stopping rules.

supply to care for them. As the children get older, male and female children contribute to or detract from their parents' income. I investigate the theory with Indian data and so will put the paper in an Indian context. In India, aside from any labor income children accrue, boys bring in dowries when they marry, while parents must pay dowries and wedding costs to get their daughters married. These dowries can be large. For example, Rao (1993) estimates that dowries can be as much as 50% of a household's assets. Furthermore, the prevalence of the joint household means that having a son creates a future expectation of more household workers, namely the son's future wife and their children. A daughter on the other hand leaves with her dowry and labor supply and can no longer be expected to contribute to her parents' household. Sons provide income security in old age, daughters do not. Even if daughters could help their parents in their old age, women suffer from having much lower income prospects than men. Sons and daughters have opposite future income effects on their parents and these differences are likely to cause differences in childhood health investment. In the poor households of India, low child health investment is likely to substantially increase child mortality rates.

The first model formalizes fertility and investment decisions using a two-period household model where the husband and wife act as a unitary utility maximizing agent. In the first period, parents decide how many children to have and how much to invest in their health in childhood. The proportion of children that survive depends on the level of investment. In the second period, parents marry off their children. A boy stays in the household, adding his income and his wife's income and dowry to the household, as well as other possible benefits such as retirement security. Girls leave their parents, taking with them their labor supply and dowry. The difference between boys' and girls' future benefits and costs causes a tension for parents between wanting to keep their children alive and parents' desire to consume resources for themselves. This tension drives the theoretical results. Given a fixed number of siblings, girls are better off with brothers and boys are better off with sisters. Intuitively, boys mitigate the costs of their sisters allowing parents to keep their girls alive. If a boy has many sisters, parents have a stronger incentive to keep the boy alive to help pay for the costs of the girls. Thus, we should expect households with a higher proportion of daughters to have higher female mortality rates and lower male mortality rates.

The empirical approach to testing this theory has two steps. First, using a large Indian demographic and health survey, I show that the sex of the first-born child is random, and, in particular, that parents are not likely to selectively abort their first pregnancy. I discuss a number of possible biases in the data and their implications. Second, a reduced form approach uses the sex outcome of the first pregnancy as a natural experiment. A household that has a first-born boy has fewer children and a higher proportion of boys than a household with a first-born girl. A first-born boy causes boys to have higher mortality rates, while causing girls to have lower mortality rates. That the mortality rates of boys actually go up in smaller, boy-proportioned households means that the sex composition effect is stronger than the household size effect for boys. Several robustness checks ensure the validity of the results. I find that the outcome of the first birth can explain about a quarter of the child mortality gap between boys and girls. The results also indicate that sex-selective abortion will reduce the child mortality gap by improving girl mortality and worsening boy mortality. Evidence is provided that the sex outcome of the first-birth affects vaccinations of higher order births, supporting the idea that fertility decisions are exacerbating parents discrimination in their provision of health resources.

The last part of my paper investigates dowries and inheritance rights that favor sons, two of the possible institutions in India that create the economic incentives for discrimination against girls. First, I examine whether dowries can explain why girls are costly. Using a detailed data set of economic information on rural Indian households, I find that while dowries are a burden for parents, and may contribute to discrimination against daughters, they are less likely to explain the benefits of sons. In particular, I find that it is plausible that women are a net economic burden to a household. Thus, a dowry may only help to mitigate the cost of a daughter-in-law. However, it is clear that the elimination of dowries will reduce the benefit of having a son and reduce the cost of having a daughter.

Second, I examine the implications of a recent amendment to the Indian constitution that gives women inheritance rights over their parents' land. I model a household which desires to maximize its bequest per son. This model leads to fertility and child mortality outcomes that are consistent with my empirical results on family formation and child mortality. The model predicts that the granting of inheritance rights to women causes parents to decrease their health investments in both their sons and daughters, raising child mortality rates. Several states granted women inheritance rights before the constitutional amendment. I use these law changes in a difference-in-difference analysis to test whether inheritance rights affect child mortality. I find that granting women inheritance rights has the unintended consequence of raising child mortality, regardless of sex.

2 Background

Ever since Visaria's book on the sex ratio of India, demographers and economists have been concerned with the relatively low numbers of women in India compared to other parts of the world. Visaria performed a detailed analysis of data available up to the Indian Census of 1961 and concluded that the low number of women can be completely explained by differences in child mortality, rather than differences in sex ratios at birth or other possible explanations. Sen (1990) was the first economist to articulate the plight of "missing women", estimating that in 1990, 100 million more women would have been alive if given the same health and nutritional resources as males. Refinements by Coale (1991) and Klasen (1994) provided slightly smaller but still large estimates of the number of missing women. Recent estimates of sex-selective abortions in India are high (Jha et al., 2006). However, Anderson and Ray (2008) look into the causes of missing women and find, like Visaria, that most of India's missing women come from excess female mortality, rather than selective abortion or infanticide.

Several studies have documented the relatively poor treatment of girls. Chen et al. (1981) find that girls in Bangladesh are given less food and are less frequently given medical treatment than boys. However, the consensus is that differences in the allocation of health care, rather than nutrition, are likely to blame for excess female mortality in South Asia (Basu, 1989; Hazarika, 2000; Asfaw et al., 2007). Apart from the evidence for direct discrimination against girls, there has been little research on the economic incentives for parents to discriminate. Rosenzweig and Schultz (1982) provide an economic argument that gender discrimination is caused by the relative income of males versus females. They provide evidence of this in India and more recently Qian (2008) finds evidence of the importance of labor income for sex differences in mortality in China.

One of the growing costs of daughters, even with rising prosperity in India, is the rise of dowries (Anderson, 2003). Anderson suggests that 93-94% of marriages in India include a dowry payment and that these payments can amount to as much as six times a household's annual income. Since the labor income of women is presumably becoming more valuable as India develops, the gain to a household of a productive daughter-in-law and her dowry income may be substantial. I provide some new evidence suggesting that a daugher-in-law, at least in rural India, may still be a net economic burden to a houeshold, even with her dowry. In this case, while dowries are a burden to parents of daughters, they are not a boon to parents of sons. Another peculiar institution of India is inheritance rights that favor sons. Inheritance institutions have not been looked at as a cause of child mortality, although their effects have been examined on economic growth (Foster and Rosenzweig, 2002). This is the first paper to argue that if parents have a strong preference to maximize the amount of land bequeathed to each son, then granting inheritance rights to women will create perverse incentives for parents to lower their health investments in both their sons and daughters.

Scrimshaw (1978) introduced the idea that parents may be purposefully causing infant mortality in order to regulate family size. Das Gupta (1987) and Muhuri and Preston (1991) followed up on that idea by examining the effects of sibling sex composition on child mortality in India and Bangladesh respectively. They both find that girls appear to have lower mortality rates if they have brothers and boys appear to have lower mortality rates if they have sisters. Pande (2003) finds that boys who have older sisters and girls who have older brothers are more likely to be immunized and avoid stunting, although she attributes this to a desire for gender balance within the household.

Cigno (1998) develops a theoretical fertility model that explicitly endogenizes childhood survival. However, he does not distinguish between boys and girls. Rosenzweig and Schultz (1982) provide an economic model linking the future income of boys versus girls in India to childhood survival. My model combines these models by both treating fertility decisions as endogenous and allowing boys and girls to have different future benefits and costs for parents. My paper is also an extension of Becker and Lewis (1973)'s quality/quantity trade-off in having children, just treating boys and girls separately.

The pioneering work of Deaton (1989) is an exception to the literature that finds discrimination against girls in South Asia. One problem with Deaton's approach is that he treats family structure as if it were exogenous. Garg and Morduch (1998) investigate the effects of sex composition on child health in Ghana. They get around the endogeneity problem in Ghana because parents there appear to not care about the sex of their children in making fertility decisions. They find that all children are better off with sisters, while I find that having many sisters is only good for sons. The difference in findings is no doubt due to the differences in contexts between Ghana and India. I resolve the endogeneity issue in India by using the sex outcome of the first birth to jointly predict number of children and the sex composition of these children. Dahl and Moretti (2007) use a similar estimation technique in the U.S. and find that first-born girls are disadvantaged compared to first-born boys. For example, first-born girls are less likely to live with their father and parents who have a first-born girl are more likely to be divorced. Interestingly, a first-born girl in the U.S. predicts higher fertility, although by only about one-fiftieth as much as in India.

3 Theory

I present a two-period model of fertility and child mortality. In the first period, parents make fertility decisions conditional on previous birth outcomes, then decide how much health capital, e.g. food and medical care, to invest in each child. There is a fixed cost to having each child, regardless of how much the parents invest, e.g. a reduction in mother's labor supply in having and caring for a baby. At the end of the first period, children die via a survival function, where the more parents invest in child health, the fewer children die. For simplicity, boys and girls are assumed to have the same survival function.[3] In the second period, the children become adults. Parents suffer a fixed cost for each daughter and parents receive a fixed benefit for each son. Thus, parents lose income if they have more surviving daughters than sons. One can think of a girls' cost being her dowry at the time of marriage, while a son's benefit is his labor income as he remains in the joint household and possibly the labor supplied by his new wife and children as well as his dowry income. The timing of the model is illustrated below:

[3]This assumption does not have an effect on the comparative statics of the model.

- Period 1.
  - 1. Fertility decisions.
  - 2. Health investment in children.
  - 3. Child mortality occurs.
- Period 2.
  - Parents pay the cost of surviving daughters and collect the benefits of surviving sons.

The parents act as a unitary utility maximizer.[4] Parents first make the decision to have children, where they either continue to have a child or stop fertility altogether, conditional on previous fertility outcomes. After fertility has stopped, parents invest in these children. Parents use backwards induction in making their decisions: they make their fertility decisions based on how their investments and, hence, expected lifetime utility, are expected to change if they have an additional child. Parents care about their own consumption in each period, c_j ($j = 1, 2$). For simplicity, I assume that given N children, a continuous proportion π of them are boys, and $1 - \pi$ are girls, where $0 \leq \pi \leq 1$. Parents also care about the proportion of children who survive, $p(k_i)$, where the proportion of children of sex i surviving is a positive, strictly concave function of the average health capital invested in children of sex i, k_i, and $0 \leq p(k_i) \leq 1$. For simplicity, parents are assumed to know the exact proportion of children who survive given the health investment.[5] Thus, if parents invest k_B in their boys, then $p(k_B)\pi N$ boys will survive to adulthood. Parents with N children, and who have decided to have no more children, face the following lifetime utility function:

$$U_T = U_1(c_1) + U_2(c_2) + A(p(k_B)\pi N + p(k_G)(1 - \pi)N) \tag{1}$$

The parents' lifetime utility (U_T) is the sum of their utility from consumption in the two periods (c_1 and c_2) and the utility of having their children survive, which, for simplicity, is assumed to be a positive constant (A) times the number of surviving children. $U_1(c_1)$ and $U_2(c_2)$ are assumed to

[4] See Eswaran (2002) for a model of fertility and mortality that includes intrahousehold bargaining.

[5] I follow along the lines of Rosenzweig and Schultz (1982). I avoid the complexity of probability distributions that are in Cigno (1998) and discrete children with binomial survival distributions as in Sah (1991). Note that the analytic results could change if expected utility and probability distributions of child survival are used, depending upon the choice of utility function, distribution, and risk aversion parameters.

be positive and strictly concave with respect to consumption. I assume that parents care about the survival of each child equally and, in the absence of their desire to spend on themselves, would equally allocate all their resources to their children. This assumption about survival utility highlights the tension in parents' allocation decisions: they want their children to survive, but they also want to consume resources for themselves. It may be the case that in reality parents care more intrinsically about a son surviving than a daughter or vice versa and this is what drives discrimination. However, the model shows that discrimination will follow from economic incentives, even without different intrinsic preferences over child survival by sex.

Parents have budget constraints in each period. In the first and second periods, parents receive exogenous incomes of Y_1 and Y_2 respectively. There is a fixed cost of F per child in the first period. This fixed cost is part of the household size effect: the more children there are the less resources there are. The other aspect of the household size effect is also built into the model via the fixed income of parents: the more children there are the less resources there are per child. This is the household size effect: an extra child both reduces the available resources and the available resources per child. In the second period, if children survive, parents must pay for daughters, but benefit from sons. For simplicity, the future cost or benefit of each daughter or son is fixed at a positive number D. I.e. it is assumed that D is unaffected by early childhood health investments. Households cannot save, borrow, or accumulate assets. This assumption about credit and saving constraints is key to Proposition 1 below and is further discussed in Appendix B.[6] A household with N children faces the following constraints for each period:

Period 1 budget constraint: $c_1 + NF + \pi N k_B + (1-\pi) N k_G \leq Y_1$

Period 2 budget constraint: $c_2 \leq Y_2 + \pi N p(k_B) D - (1-\pi) N p(k_G) D$

In the first period, parents spend their income on themselves, the fixed costs of having N children, and any investments they wish to make in their children. In the second period, parents consume whatever is left over of their income net of the costs of their surviving children. Parents make all their decisions in the first period, i.e. how many children to have and how much to invest

[6] As shown in Appendix B, if there are perfect credit markets, I predict the opposite result of Proposition 1. However, from the empirical results below, we can see that Proposition 1 must hold, otherwise the coefficient on γ_{boy} would be the opposite sign.

in each child.

From the model, the following three propositions hold given that D is sufficiently large (proofs given in Appendix A).

Proposition 1. *Assume fertility decisions have stopped (i.e. given a fixed N). The greater the proportion of boys in a family, the less is invested in each boy:* $\frac{\partial k_B}{\partial \pi} < 0$.

The result for Proposition 1 comes from the parents' desire to transfer their future income gains from additional sons' future income to the present. They do this by reducing their expenditure on boys in childhood. Thinking about what happens if the proportion of boys goes down is more intuitive. Imagine a boy with many sisters. Parents face a large future cost from their daughters, and so, to help to reduce the future burden, they will want to ensure that their son survives, investing more in the health of that son than if he had brothers.

Proposition 2. *Assume fertility decisions have stopped. The greater proportion of boys in a family, the more is invested in each girl:* $\frac{\partial k_G}{\partial \pi} > 0$

Proposition 2 comes from the overall income gains from additional sons, which allow parents to spend a bit more on girls. Imagine a girl with many brothers. The future costs of that girl are ameliorated by the presence of sons, and thus parents can better afford to keep the girl alive.

Proposition 3. *Assume that there is a 50% probability of having a boy or girl, assume Household 1 (HH1) has relatively more boys and Household 2 (HH2) has relatively more girls, (i.e.* $\pi_{HH1} > \pi_{HH2}$*), and both households have N total children, then parents in HH2 have a larger utility gain from increasing N than parents in HH1. That is, parents with relatively more girls have a stronger incentive to continue having children.*

Proposition 3 follows from Propositions 1 and 2. A household with a high proportion of sons, compared to a household with a low proportion of sons, which then has an additional son, will invest less in each son (from Proposition 1). The high-son-proportioned household will thus expect smaller future gains from an extra son, since that son is more likely to die. The household with a high proportion of sons, compared to the household with a lower proportion of sons, which then has an additional daughter, will invest more in each daughter (from Proposition 2). The high-son-proportioned household will thus expect higher future costs from an extra daughter, since that daughter is more likely to live. Thus, the expected gain from sons is smaller and the expected loss from daughters is larger in the high-son-proportioned household, giving it on net a smaller incentive to have an extra child compared to a low-son-proportioned household. Proposition 3 is the theoretical explanation of son-preferring fertility stopping rules.

Note that a higher proportion of girls will both hurt girls via the sex composition effect and lead to higher fertility, also hurting girls. The point of the propositions is to illustrate that as the parents have children, those with girls are pushed to have more children. The resulting distribution will create a portion of households that are at a particular disadvantage to girls: Girls are in larger households than boys on average, which hurts girls, and they are in households with a high proportion of girls, which is worse than if they were in the same household with a higher proportion of boys.

4 Data and Estimation

The paper uses the 2002-2004 Reproductive and Child Health Survey (RCH II) to analyze child mortality in India. The RCH II is a nationally representative survey of approximately 500,000 women aged 13-44 who have ever been married. The survey was implemented by the Government of India via the International Institute for Population Sciences (IIPS), with the goal of better understanding demand for family planning, contraceptive use and reproductive knowledge, early child health, and utilization of health facilities. The survey was designed to be representative at the district level and, hence, the large number of surveyed women. The use of such a large dataset, as opposed to the smaller Indian National Family Health Surveys, is critical for this paper. First, it allows enough power to verify the sex ratio at birth for first-borns. Second, since only a small percentage of children die, it more precisely estimates effects on child mortality.

Ideally, if there was no endogeneity problem, it would be possible to simply regress boy and girl mortality on the number of children and their sex composition. However, higher child mortality increases the number of children parents want, which could also influence the sex composition of children. In particular, a change in the fertility rule will affect the number of children born and the sex composition of children. If families are more or less likely to selectively abort a pregnancy depending on the expected survival rates of their children, this will also alter the number and sex composition of children.

Jensen (2003) and Angrist and Evans (1998) use the outcome of the first or first two pregnancies as an instrument for the number of children. Jensen follows a similar strategy to the one followed in this paper and uses the fact that parents' preference for sons in India will cause parents with a first-born son to have fewer children than parents with a first-born girl. Angrist and Evans (1998) look at the U.S. and use parents' preference for a sex balanced family to predict more children if the first two births are of the same sex. However, Rosenzweig and Wolpin (2000) point out that for Angrist and Evans the sex composition of children can affect the outcomes of interest, and, thus, the instrument is not necessarily valid. In particular, any household size affect may be conflated with the sex composition effect. Twins as a first pregnancy are another exogenous outcome that impacts number of children born and the sex composition of these children, yet it cannot reliably be used as an extra instrument because twins are different from non-twins, in particular having lower birthweight on average, which means that, for example, a pair of boy twins are biologically weaker than two sons born separately (Rosenzweig and Zhang, 2006). To simplify the analysis below, I drop all households with first-born twins which make up approximately 0.5% of all observations.

Thus, we cannot instrument for the two endogenous variables: number of children and their sex composition. However, we can estimate the reduced form effect of a first-born boy versus first-born girl on child mortality, which in itself provides a test of household size and sex composition on child mortality. Proposition 3 above leads to the following prediction:

Prediction 1. *Parents who have a first-born girl will have a stronger incentive to continue having children than parents with a first-born son. Thus, a first-born daughter should predict more children on average, than a first-born son.*

Using the the sex outcome of the first-birth as a natural experiment, we find the direct effect of the outcome of the first pregnancy on child mortality. Note that it will be true that, even if households were not following a stopping rule and having X children of random sex, girls in a household with a first-born girl will have a higher proportion of girls on average than a household with a first-born boy. Given the household size and sex composition effects of the sex outcome of the first-birth, we can then make the following empirical predictions from the theory above:

Prediction 2. *Girls in households with a first-born boy, which predicts fewer children with a higher proportion of boys, will benefit from both the household size and sex composition effects and have higher childhood investment and, hence, lower childhood mortality than girls in households with a first-born girl.*

Prediction 3. *Boys in households with a first-born boy may face either higher or lower childhood investment. Fewer children decrease mortality via the household size effect, while the higher proportion of boys reduces mortality via the sex composition effect.*

Prediction 4. *Households have a larger incentive to use selective abortion for later pregnancies if the first-born child is a girl.*

Although selective abortion is not explicitly a part of the above model, this last prediction follows intuitively from the model: the more girls a household has, the larger the future economic burden. Thus, a household with net future losses (with many girls) would gain more from selective abortion than a household with net future gains (with many boys). For example, Jha et al. (2006) shows that the sex ratio of second-born children in India is only unnaturally skewed toward more boys if the first-born child is a girl. It is a key assumption for the empirical section that selective abortion occurs mostly in higher order births in India, so it is reassuring that the theoretical model supports this assumption. Appendix C provides evidence that the sex of the first-born child is likely exogenous in India, as opposed to higher birth orders where selective abortion may bias the outcomes. In particular, the male/female sex ratio at birth for children born within ten years of the survey is 1.066, within the range considered biologically normal (1.03-1.07), while it is significantly higher for later births. Since parents are not selectively aborting their first pregnancy, the sex outcome of the first pregnancy will actually tell us what would have happened to a household if it had used selective abortion for the first pregnancy. Ignoring the direct costs of selective abortion, if we ask what households with first-born girls would look like if they had selectively aborted the first-born girls, the answer is that, on average, they would look like families with first-born boys. Thus, the sex outcome of the first pregnancy is also a test of the impact of sex selective abortion on child mortality.

The outcome of the first pregnancy strongly predicts the number of children born. Table 1 shows a regression of the number of children born on the outcome of the first pregnancy (FirstBornBoy = 1 if the first-born is a boy, 0 if a girl) along with some additional independent variables, which will be explained below. The estimation equation is as follows:

$$\#ChildrenBorn_{ij} = \gamma FirstBornBoy + \beta_1 X_i + \beta_j State_j + e_{ij} \tag{2}$$

The number of children born in household i and state j depends on the outcome of the first pregnancy, household characteristics (X_i), and a state dummy variable ($State_j$). A first-born son predicts a decrease in number of children by approximately one quarter of a child and more than one third of a child if we restrict the sample to women who are likely to have completed their fertility: women over 35 or women who are sterilized.[7] Not surprisingly, a first-born son strongly predicts about a third higher proportion of sons in a household.

The paper focuses on the child mortality of boys or girls, separately, of birth order two and higher. The dependent variable, child mortality, is defined as the percent of children who die between the ages of one and sixty months, multiplied by 100. The model yields predictions on average boy and average girl mortality instead of on specific birth orders, which is why the empirical results will focus on average mortality. Average mortality is also useful in that it allows us to ignore the number of children born, which is endogenous. Children who die before one month are ignored because most deaths at this age are from birth defects or other issues not within parents' control.[8] Sixty months is a cut-off used in most of the literature on child mortality because a high proportion of child deaths occur before age five. First-born children are more likely to die than later born children (Hobcraft et al., 1985). Thus, it would not be surprising to find that if we include first-

[7] To put these numbers in perspective, Angrist and Evans (1998) only predicted a fertility increase of 0.06 children if the first two children were of the same sex, while Dahl and Moretti (2007) predicts an increase of 0.007 children if the first-born is a girl in the U.S.

[8] See, for example, Simmons et al. (1982), Simmons et al. (1978), and Smucker et al. (1980). The results in Table 3 are robust to including these deaths in the specification.

Table 1: OLS: Dependent Variable = Total Children Born

Variable	**All Women**	**Mother Age $\geq$ 35**	**Mother Sterilized**
FirstBornBoy	-0.241^{**}	-0.356^{**}	-0.412^{**}
	(0.005)	(0.010)	(0.009)
Age Mother	0.125^{**}	0.068^{**}	0.066^{**}
	(0.002)	(0.002)	(0.001)
Education Mother	-0.067^{**}	-0.081^{**}	-0.057^{**}
	(0.001)	(0.002)	(0.001)
Education Father	-0.033^{**}	-0.043^{**}	-0.028^{**}
	(0.001)	(0.001)	(0.001)
Hindu	0.033^{*}	-0.022	$0.031^{\dagger}$
	(0.014)	(0.024)	(0.018)
Muslim	0.675^{**}	1.052^{**}	0.659^{**}
	(0.019)	(0.035)	(0.025)
Scheduled Caste	0.351^{**}	0.480^{**}	0.417^{**}
	(0.008)	(0.017)	(0.012)
Scheduled Tribe	0.304^{**}	0.352^{**}	0.301^{**}
	(0.013)	(0.024)	(0.018)
Backwards Caste	0.132^{**}	0.167^{**}	0.041^{**}
	(0.007)	(0.013)	(0.010)
Rural	0.143^{**}	0.162^{**}	0.041^{**}
	(0.008)	(0.013)	(0.009)
Constant	-1.230^{**}	0.924^{**}	0.069^{**}
	(0.096)	(0.177)	(0.010)
State Dummies	Yes	Yes	Yes
N	443944	150782	156942
Clusters	1590	1590	1559
R^2	0.418	0.315	0.307
F	1002.87	669.67	737.39

Significance levels : $\dagger$: 10% $*$: 5% $**$: 1%

Notes: Errors clustered by primary sampling unit in parentheses, no households with first-born twins.

births, boy mortality is higher and girl mortality is lower amongst households with a first-born boy. Given this fact, the mortality estimates are only carried out for children of birth order two and higher. Because the paper focuses on fertility decisions, the paper only provides estimates for households where the mother is aged 35 and older, when she is likely to have completed her fertility. Only about 12% of women in the RCH II have a child at age 35 or older, and more than 70% of women aged 35 and older have been sterilized. However, as explained in Appendix C in more detail, there is evidence of recall and survival bias in this age group, whereas there is little if any such bias for younger women.

Thus, there is a trade-off between fertility completion and bias. Since the paper focuses on fertility, I choose fertility completeness and acknowledge the potential bias. Recall bias occurs when parents either misreport or do not remember the true outcome of their first pregnancy. This might happen, for example, if women systematically do not report first-born girls who die in infancy. Such recall bias has been reported in China (Smith, 1994) and in India for the National Family Health Surveys (IIPS, 1995). Furthermore, since a first-born girl predicts higher fertility, and maternal mortality in India is relatively high, we may expect there to be fewer surviving mothers at the time of the survey who had a first-born girl many years in the past.[9] I call this effect on mother mortality "survival bias". In either case, the bias will cause first-born boy households to appear to have higher mortality rates than first-born girl households. Parents that do not report a first-born girl who died in infancy are likely to be worse-off and suffer higher child mortality rates than other households. Mothers who die from having more children because the first-born child is a girl are more likely to be in worse-off households with higher child mortality.

Mother's age, parents' education in years[10], dummy variables for state, caste, religion, and whether the household is in a rural or urban location are chosen as independent variables. These are all variables which could affect child mortality and are exogenous to child survival. Table 2 presents summary statistics of these variables and shows any differences between first-born boy

[9] The Indian National Family Health Survey 1992/93 estimates that approximately 0.5% of pregnancies end in the mother's death. Also see http://unstats.un.org/unsd/demographic/products/indwm/tab3b.htm for maternal mortality statistics.

[10] Illiterate individuals are coded as having no years of education, which is not necessarily true. The results are robust to simply including dummy variables for literate/illiterate instead of years of education.

and first-born girl households. On average, first-born boy households have parents who are less educated than parents in first-born girl households. These differences are in accord with recall and survival bias, which make first-born boy households appear worse off than they would be without these biases. Mothers are older in first-born boy households, which is consistent with survival bias because women who had first-born sons are likely to live longer. However, although statistically significant, these differences are small and indicate that the bias in the data is not large.[11]

Even though I show in Appendix C that selective abortion is unlikely for first pregnancies, it is useful to discuss its implications for my estimates. One problem is that if selective abortion were occurring, it is unclear what the bias would be. On the one hand, better-off households are the ones with access to the health resources and wealth required to undergo a selective abortion. This would cause boy-first households to appear better off. On the other hand, the worse-off households are the ones who may benefit most from this practice and have the largest incentive to selectively abort. This would cause the opposite bias, making boy-first households appear worse. If parents who resort to selective abortion simply hate girls and love boys, the bias would be toward higher girl mortality and lower boy mortality in boy-first households. Hypothetically, if selective abortion was responsible for some of the skewed sex ratio of first-borns, the summary statistics indicate that on average, worse-off households are resorting to selective abortion of the first pregnancy. Again, the bias would be in the direction of there being higher child mortality amongst first-born boy households.

Given that first-birth outcomes are exogenous to child mortality, we can estimate the following equation:

$$M_{S,ij} = \gamma_S FirstBornBoy + \beta_1 X_i + \beta_j State_j + e_{ij} \tag{3}$$

where $M_{S,ij}$ is the average child mortality (as explained above) of children of birth order two and higher of sex S in household i in state j. X_i is a set of household characteristics: parents' ages and

[11] We may expect a similar bias if infanticide was responsible for the above normal sex ratios of first-births in older women, since we would expect only the families with the worst socioeconomic situation to resort to such measures. Note that a first-born boy predicts an approximately extra three weeks between the first and second birth. There is evidence that shorter birth intervals cause low birth weight and, hence, higher mortality rates (Gribble, 1993). This would bias the results in the opposite direction and result in higher mortality rates amongst first-born girl households.

Table 2: Independent Variables by First-Birth Outcome, Wife Age ≥ 35

Variable	Entire Sample	First-Born Boy	First-Born Girl	Difference
Mother Age (years)	39.005	39.019	38.989	0.031*
Mother Education (years)	3.788	3.755	3.825	-0.070*
Father Education (years)	6.538	6.490	6.571	-0.082*
Hindu (0/1)	0.751	0.749	0.753	-0.005*
Muslim (0/1)	0.109	0.112	0.105	0.007*
Scheduled Caste (0/1)	0.154	0.155	0.154	0.001
Scheduled Tribe (0/1)	0.162	0.162	0.162	0.000
Backwards Caste (0/1)	0.359	0.358	0.261	0.003
Rural (0/1)	0.649	0.649	0.649	0.001
N	152104	80714	71390	

Notes: No households with first-born twins. N is slightly smaller for mother and father education due to missing values.
* indicates significant at the 5% level using t-tests for age and education, and Pearson chi-squared tests for the other variables.

education, urban/rural, caste, and religion. $State_j$ is a dummy variable for state. $FirstBornBoy$ is a dummy variable which takes on the value of 1 if the first-born child in a household is a boy. The predictions are that $\gamma_{girl} < 0$, since a first-born boy benefits girls through the household size and sex composition effects, while $\gamma_{boy} > 0$ if the sex composition effect is stronger than the household size effect, and $\gamma_{boy} \leq 0$ otherwise. As mentioned above, recall and survival bias will cause first-born boy households to appear worse off than they really are, raising the coefficient on $FirstBornBoy$. Thus, any positive coefficient is suspect in the sense that its sign may be due to bias.

4.1 Estimates

Table 3 shows the impact of a first-born boy on the percentage of siblings who died during childhood. A first-born boy lowers average girl mortality by just over 0.3%. Note that the biases mentioned above will tend to push this coefficient towards a positive number, in this case making the coefficient smaller than it should be. Since the coefficient is negative, we can safely conclude that girls with a first-born older brother, with fewer siblings and a higher proportion of boys, have lower mortality rates than girls with a first-born sister. Fertility decisions in India, via the approximate use of stopping rules, cause girls on average to be in households with more children, and in households with a higher proportion are girls. These fertility decisions translate into less resources for girls (household size effect) and increased discrimination (sex composition effect) and, hence, higher mortality rates amongst girls.

Table 3: OLS: Dependent Variable = Percent of Boys or Girls Dead × 100 ⋆

Mother Age ≥ 35, Birth Order ≥ 2		
Variable	**Boys**	**Girls**
FirstBornBoy	0.471**	-0.313**
	(0.087)	(0.109)
Age Mother	0.046**	0.115**
	(0.016)	(0.019)
Education Mother	-0.112**	-0.190**
	(0.012)	(0.016)
Education Father	-0.169**	-0.244**
	(0.011)	(0.016)
Rural	0.452**	1.163**
	(0.105)	(0.138)
Constant	5.007**	0.342
	(1.717)	(1.483)
Religion Dummies	Yes	Yes
Caste Dummies	Yes	Yes
State Dummies	Yes	Yes
N	117734	104494
Clusters	1587	1588
R^2	0.025	0.040
F	55.81	76.35

Significance levels : † : 10% ∗ : 5% ∗∗ : 1%

Notes: Errors clustered by primary sampling units in parentheses, no first-born twins households.

⋆ Percentage of children born at least 60 months before survey and died between the ages of 1 and 60 months.

Boys are about 0.5% less likely to survive in households with a first-born boy. It is possible that bias in the data has caused this positive coefficient. However, as shown above, the bias is likely small. In addition, although I do not show the estimate here, if the sample is restricted to women under 35, with less bias, the coefficient remains positive and significant, providing further support that the bias is unlikely the cause of the positive value. Thus, if bias is not strong enough to account for the entirety of the positive coefficent, the results indicate that the sex composition effect is stronger than the household size effect for boys. This result provides evidence that the need is so strong in the first-born girl households to keep boys alive that boys are given a larger share of the resources, enough to outweigh any drop in resources from having extra siblings. The result indicates that sex composition can have a strong effect on child survival outcomes. These results also point out that discrimination is not just anti-girl, it is also pro-boy. Thus, the reason that girls have higher childhood mortality rates than boys is not just that parents are more likely to neglect girls in larger, girl-proportioned households. In these households, parents actively improve the health of boys, making them better off than if the boys had fewer sisters.

The results also provide insight into what happens when parents use selective abortion. If parents had used selective abortion for their first pregnancy, i.e. they had first-born boys instead of first-born girls, the estimates predict that non-aborted girls would be better off, but boys would be worse off. Given the above coefficients, the result could be net higher child mortality, even if it would lower the gap between boy and girl child mortality. This result is in contrast to Lin et al. (2008) who find that both boys and girls had lower mortality rates soon after selective abortion became legal in Taiwan in the mid-1980's, although the magnitude of these effects are much smaller than I find, since mortality rates were already very low in Taiwan.

I estimate the amount of the child mortality gap that the outcome of the first pregnancy accounts for, restricting my sample to women aged 35 and older and children of birth order two and higher. Approximately 49 out of 1000 boys die between 1 and 60 months, while approximately 69 out of 1000 girls die. Thus, girls suffer about 40% higher child mortality than boys for this group. If we unconditionally restrict ourselves to households with a first-born boy, the gap closes by about a third, so that girls only suffer about 28% higher child mortality than boys. Using the estimates from Table 3 to control for the other independent variables, if all children were in first-born boy households (taking all girl-first households and subtracting γ_{boy} and γ_{girl} from boy and girl mortality respectively), the gap would close by about a quarter to slightly under 31%, which represents a large reduction in the child mortality gap.

4.2 Robustness Check: Logit Analysis

Average child mortality has discrete steps in it, and in particular values will be grouped at, for example 0%, 25%, 33%, etc. Thus, although the theoretical model is about average child mortality, one may object that OLS is not the correct estimation model. To verify the above results, I run a logit estimation, where the estimation equation is the same as in Equation (3), except that the outcome, $M_{S,ij}$, is now a 0 or 1 variable, which is 1 if the second-born child of sex S died during childhood. I only use the second-born child because almost everyone has at least two children. By the age of 35, 95% of women have two or more children, and 99% of sterilized women have at least two children. Thus, we should not worry much about parents not selecting to have that second child. However, we should worry about higher order births, since many parents will not have more than two children. The results of the logit estimation are shown in Table 4.

Table 4: LOGIT: Dependent Variable = 2nd Order Boy or Girl Dead (0/1)

Mother Age $\geq$ 35		
Variable	**Boy**	**Girl**
FirstBornBoy	0.089*	-0.065†
	(0.036)	(0.035)
Age Mother	0.033**	0.044**
	(0.007)	(0.006)
Education Mother	-0.074**	-0.084**
	(0.007)	(0.007)
Education Father	-0.040**	-0.041**
	(0.005)	(0.004)
Rural	0.186**	0.173**
	(0.048)	(0.044)
Constant	-3.386**	-5.07
	(0.523)	(0.781)
Religion Dummies	Yes	Yes
Caste Dummies	Yes	Yes
State Dummies	Yes	Yes
N	75062	68525
Clusters	1544	1540
Pseudo R^2	0.066	0.070

Significance levels : $\dagger$: 10% $*$: 5% $**$: 1%

Notes: Errors clustered by primary sampling units in parentheses, no first-born twins households.

Calculating the marginal effect of a first-born boy on the mean household I find that a first-born boy increases the probability of a second-born boy dying by 0.27% and decreases the probability of a second-boy girl dying by 0.24%. These findings are similar to those in the OLS estimation.

4.3 Further Analysis of Results

A lifetable is provided in order to better understand the empirical results. The table shows how childhood survival varies by sex and by the outcome of the first birth. As above, the survival estimates are restricted to children of birth order two and higher, so as not to conflate the additional mortality a first-born may suffer compared to higher order births with the impact of the sex of the first-born on the mortality of other siblings. The results are shown in Figure 1. Blue lines represent boy survival, while red lines represent girl survival.[12] Solid lines are households in which there is a first-born boy (BF) and dotted lines are households with a first-born girl (GF). Each age point shows the proportion of all children who were born at least that many months before the survey was taken who survived up until that age.

There are several salient features of this figure. First, the survival lines are in accord with the empirical results above: A first-born boy causes girls to have a higher chance of survival while causing a boy to have a lower chance of survival. Second, the survival gap between boys is larger than the gap between girls. The relatively large differences in survival rates between boys of the different household types compared to girls may be due to recall/survival bias as stated above. In this case, children in boy-first households would seem to be worse off than they really are because the worst girl-first households are either missing or are misrecorded as boy-first. Third, although it is commonly believed that boys suffer higher mortality rates in the first year of life than girls (Hill and Upchurch, 1995), in girl-first households boys survive more than girls starting at six months of age. Fourth, the survival of girls in first-born boy households after age 24 months is very close to the survival of boys in first-born boy households. The conclusion from the figure is striking: a large portion of the mortality gap between boys and girls would be eliminated if children were born into boy-first households. At age 60 months, the survival gap between boys and girls is about 22.3% for girl-first households and only 1.3% for boy-first households. If there were no bias against girls we would see significantly higher boy mortality than girl mortality. So that even though the gap for boy-first households seems small, there is still significant bias against girls in boy-first households.

[12] For those reading a black and white copy, the blue lines appear a darker grey than the red lines.

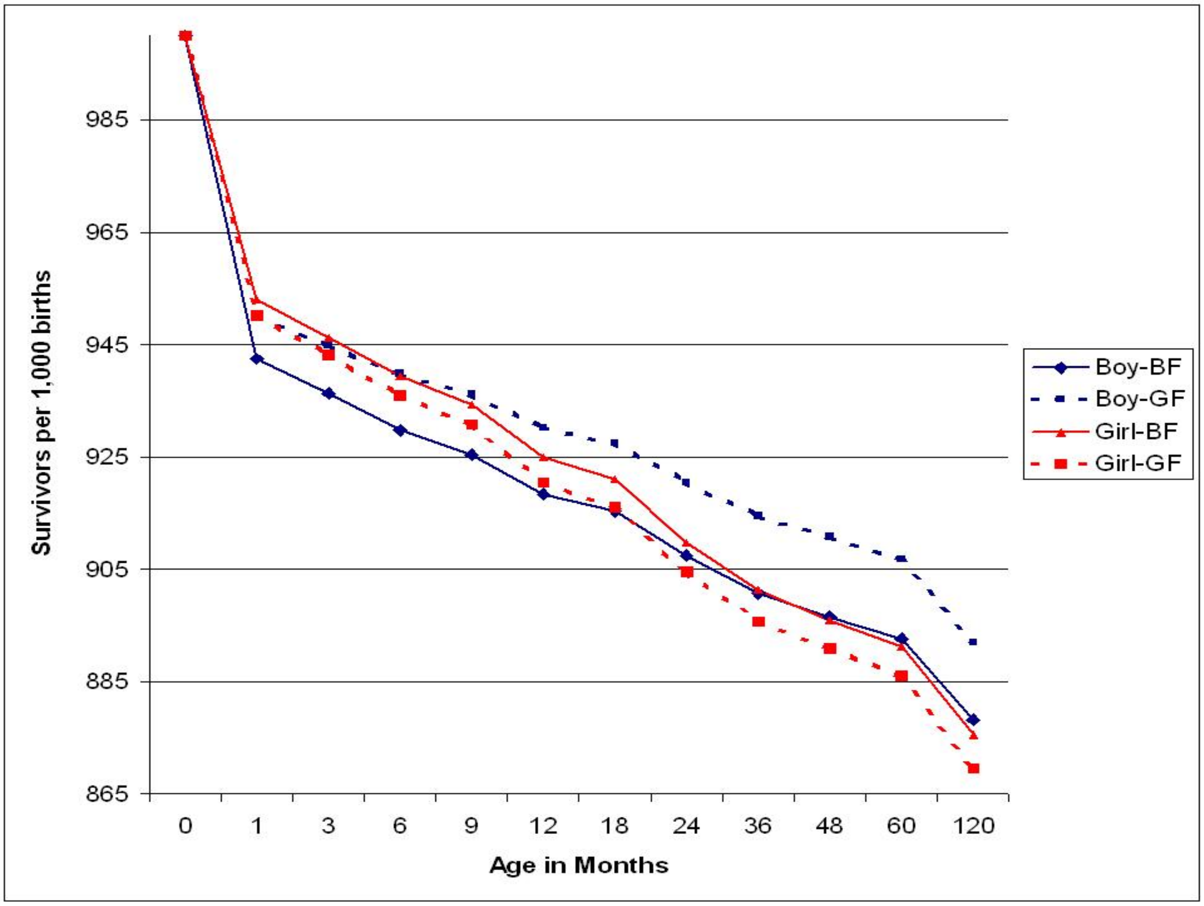

Figure 1: Number of Surviving Children per 1,000 births, 2nd Order and Higher, by Sex and Outcome of the First Pregnancy (BF=First-Born Boy, GF=First-Born Girl), No Households with First-Born Twins, RCH II, All India, Sample Weights Used

4.4 Vaccinations as Evidence of Health Resource Discrimination

There may be several mechanisms through which the outcome of the first pregnancy leads to differential child mortality. For example, sibling rivalry may account for some of the mortality results. Older sisters may protect and care for younger brothers but not do so for younger sisters. Older brothers may protect and care for their younger sisters, but not their younger brothers. This paper has not determined the exact mechanism through which the first pregnancy affects the well-being of future children, but it is clear that some causal process exists between fertility decisions and child mortality outcomes.

Examining the effect of the first pregnancy on health provision instead of directly on mortality allows us to understand the form that child health resource discrimination takes. Such an estimation may provide evidence that family formation leads to health resource discrimination by parents as opposed to some other mechanism through which mortality could occur. The RCH II asks mothers about the vaccine status of their most recent one or two births after January 1, 1999 or January 1, 2001, depending on the phase of the survey.[13] Whether siblings are vaccinated is not likely an independent event. Thus to avoid this issue, the sample is restricted to only the single most recent birth of a mother. About 76% of recently born boys have been vaccinated, while about 72% of recently born girls have been vaccinated.[14] This vaccination gap is not large, yet nonetheless could be one of the causes of higher mortality amongst girls. Oster (2008), for example, shows that vaccinations can account for as much as 20-30% of excess female mortality in India.

The effect on vaccinations is estimated in two ways: First, whether the child received any vaccinations and, second, the number of vaccinations received. The first estimation uses the following linear probability equation:[15]

$$PV_{S,ij} = \gamma_{S,PV} FirstBornBoy + \beta_1 X_i + \beta_j State_j + e_{ij} \tag{4}$$

where $PV_{S,ij}$ is a 0/1 variable equal to 1 if child i in state j of sex S was vaccinated, for children of birth order two and higher. As with mortality, $PV_{S,ij}$ is multiplied by 100 to make the results

[13] About 20 children were included even though they were born before the cutoff dates.

[14] Similar discrimination against girls in vaccinations is reported in Borooah (2004).

[15] The estimation is robust to logit and probit specifications, yielding similar results.

easier to read. X_i is as above, although child age in months is added as an independent variable. If vaccinations are one route of discrimination of health resources, we expect to see $\gamma_{B,PV} < 0$ and $\gamma_{G,PV} > 0$. I.e. if a boy is born first, this should cause later born boys to get vaccinated less and girls to get vaccinated more. A similar estimation is carried out with number of vaccinations as the dependent variable. Because so few women aged 35 and older have a recent birth, restricting ourselves to this sample will mean large standard errors. Thus, I present results for mothers at least age 35 and all mothers. The coefficients on FirstBornBoy are shown in Table 5.

Table 5: OLS:Vaccinations, Coefficient on FirstBornBoy

Most Recently Born Child, Birth Order $\geq$ 2				
	Dependent Variable			
	Vaccinated (0/1) * 100		**# Vaccinations**	
	Boys	**Girls**	**Boys**	**Girls**
Mother Age $\geq$ 35	-1.345	0.819	-0.134*	0.095
	(0.937)	(0.993)	(0.066)	(0.070)
N	8046	7203	8046	7202
All Ages	-1.592**	0.268	-0.141**	0.023
	(0.274)	(0.313)	(0.021)	(0.024)
N	73621	63806	73612	63804

Significance levels : † : 10% $*$: 5% $**$: 1%

Notes: Errors clustered by primary sampling unit in parentheses, coefficients on Child Age, Mother's Age, Parent's Education, Religion, Caste, Rural/Urban, and State Dummy Variables not shown.

In accord with the mortality estimates above, a first-born boy predicts about 0.14 fewer vaccinations and a 1.3-1.6% smaller probability of being vaccinated for boys, although the result for older mothers for the probability of getting vaccinated is not significant. The estimations predict between 0.02 and 0.1 more vaccinations and between a 0.3 and 0.8% larger probability of being vaccinated for girls, although the results for girls are not statistically significant. Thus, there is evidence that fertility decisions lead to discrimination in health resources by parents, at least amongst boys.

5 Economic Incentives: Dowries and Inheritance

Although the theoretical model above uses the fact that boys are economically beneficial to parents in the future while girls are economically costly, I have not shown empirically why this is so. Below

I present data on dowries, which show that they can play a part in the relative costs and benefits of boys versus girls. I will also show a complementary model to the cost/benefit one. In this model, parents care about maximizing their bequest per son, rather than their future income from their children. I provide some evidence that this alternative model has merit. In particular, the model predicts that granting women inheritance rights will give parents an incentive to reduce their health investments in their sons and daughters, and I show empirical results that help to confirm that this is so. For both my investigations of dowry and inheritance, I use the detailed economic data from the National Council of Applied Economic Research (NCAER) from their 1999 round of the Rural Economic and Demographic Survey (REDS). The data is nationally representative of Indian rural households and covers 7,474 households in 253 villages in 16 states.

5.1 Dowries

I investigate whether dowries are a large cost of having a daughter and benefit of having a son. In many cases in India there are reciprocal transfers between the families of the bride and groom at the time of marriage. I call dowry a net positive transfer from the bride's family to the groom's family, and I call brideprice a net negative transfer. In the REDS data, about 57% of heads of households report dowries at their marriage, while about 4% report brideprice, with the remaining households having zero net transfers.[16] The average net transfer to the husband's family is 8,300 Rupees (Rs.).[17] Dowries are more prevalent amongst children of the heads of household. About 73% of childrens' marriages had dowries with average net transfers of Rs. 13,200. These transfers are in the range of a quarter to a third of yearly household income, although are about double that if we restrict the sample to marriages that only involve dowries. Hence, we see that dowries are prevalent and a large fraction of household income.

Anderson (2007) points out that dowries appear to be rising in several parts of India at much faster rates than inflation. There are two main explanations for the apparent rise in dowries in India. Rao (1993) explains the rise in dowries via a change in demographics. Since men marry younger

[16] All calculations of population means are calculated using survey weights.

[17] The survey asks the heads of household to value their own and their children's marriage transfers. However, it is ambiguous as to whether they are reporting the values in 1999 rupees or in rupee values at the time of marriage.

women, with a growing population the supply of brides relative to grooms increases. Simple supply and demand would then lead us to a higher demand for grooms, causing an increase in dowries. The second explanation is due to Anderson (2003) who explains dowries via a hierarchical society that wants to maintain marriage within each social strata. As wages rise, higher social groups must demand larger dowries to prevent lower social strata from marrying up.

Another simple explanation, closer to Becker (1991), is that men's wages have been rising much more quickly than women's wages, causing women to pay more for a match in the marriage market. The underlying cost of daughters and benefits of sons may just be their relative wages as indicated in Rosenzweig and Schultz (1982) and Qian (2008) and these wage differences are what determine dowries. In particular, as Rajaraman (2005) suggests, if women contribute less than is spent on them in the household, then dowries may represent the net present cost of a woman to the household. If this is true, and daughters are always an economic burden to a household, then dowries can explain a portion of the large cost of a daughter. In this case parents always pay the full cost of their daughters. If dowries are eliminated, then the cost of daughters to their parents will drop since the cost will be transferred to the in-laws' household at marriage. From the groom's parents' perspective, dowries mitigate the cost of the new wife, but there is no large economic gain from the dowry. Of course, if women contribute to the household as much as they cost, then dowries are a large benefit from having a son, while remaining a cost from having a daughter. Regardless of whether women are an economic burden or not, if dowries are eliminated, then the economic cost of a daughter and the benefit of a son must drop, leading to more equal treatment by parents.

I cannot put a monetary value on a wife's housework or the children she provides or her non-economic contributions to the household. However, using the REDS data, I calculate that the average wages of women, including work on the family farm, are much less than per capita expenditure. Focusing on women in their main working years, between the ages of 20 and 60, I find that while mean yearly household per capita expenditure is about Rs. 6000, the mean yearly income from the wife of the head of household (or head of household if the head is a woman) is about Rs. 1800. Women spend about as much time on housework as they do on work that would be included in their incomes. If the value of housework is much more than the value of her other work or expenditure on women is much less than expenditure on men, then women may not be an economic burden to their households. However, these numbers suggest that it is not implausible that women are an economic burden and that any advantage from receiving dowries is lessened by this burden. Overall, then, dowries can help to explain why daughters represent a large future cost to parents, but are less likely to be able to explain why sons provide large economic benefits.

5.2 Inheritance Rights

Another possible explanation for discrimination against girls is the prevalence of legal inheritance rights for sons and not for daughters. Agarwal (1994) argues that women's lack of rights to agricultural land in India is a major cause of women's relative economic disadvantage. Recently Roy (2008) presents evidence that reforms that gave women improved inheritance rights over agricultural land in some states of India helped to improve women's autonomy. Although providing women greater rights over land should help women, there is a potential unintended consequence. If parents have a strong desire to bequeath their agricultural land to their sons, then granting daughters inheritance rights will cause parents to have stronger incentives to not have girls via increased selective abortion or reduction in child health investment. Giving women inheritance rights will also lead to a rise in boy mortality. Assume there are economies of scale in the amount of inherited land because there are economies of scale in farming. Then parents, wanting to maximize the amount of land given to each son, will have an incentive to have few sons. If the amount of land parents can give to sons falls, because their daughters can now inherit, then at the margin parents will want fewer sons. That is, parents may invest less in their sons' health, increasing boy mortality rates.

There are several possible reasons why parents want to maximize their bequest per son. Parents may have strategic bequest motives (Bernheim et al., 1985) because sons are able to provide greater economic support than daughters. Or there are economies of scale in farming and sons are more productive farmers than daughters, so that parents desire to maximize income per child after their death. Another reason parents may favor inheritance for sons is that daughters after marriage may be forced to effectively give their inherited land to their husband, reducing the actual benefit for daughters from the land.

To better understand the unintended consequences of giving women inheritance rights, I construct a model of fertility decisions and child health investment based on inheritance preferences. This is an alternative to the one based on different future economic costs and benefits. Assume that parents consecutively make fertility decisions and the sex outcome of each pregnancy is random. Assume parents have stopped their fertility and they have N children, of which a proportion π are sons and $(1-\pi)$ are daughters. They then have the following utility function:

$$U_T = \begin{cases} U_1(c) + Z(\frac{L}{p(k_B)\pi N + \alpha p(k_G)(1-\pi)N}) + A(p(k_B)\pi N + p(k_G)(1-\pi)N) & \text{if } p(k_B)\pi N \geq 1 \\ \\ U_1(c) - J(L) + A(p(k_B)\pi N + p(k_G)(1-\pi)N) & \text{if } p(k_B)\pi N < 1 \end{cases}$$

subject to the following budget constraint: $c \leq Y - \pi N k_B - (1-\pi)N k_G - NF$

The model ignores future economic costs or benefits of sons and daughters and parents only gain future utility from a bequest to their sons in the $Z(\cdot)$ function. For simplicity in my proofs, I assume that Z is a positive constant. That is, parents have linear utility in their bequest per son. L is the amount or value of land that parents will bequeath to their sons. α is the weight on the amount of land that girls inherit. For now assume that $\alpha = 0$, that is girls have no inheritance rights. Note that the model assumes that each son inherits equal amounts, which is in accord with the situation in India where partible inheritance for sons is the overwhelming norm. The model assumes that there are economies of scale in the size of the bequest, and parents prefer giving everything to one son. If parents have less than one surviving son, then they suffer a disutility equal to $J(L)$. I treat J as a positive constant times L. If J or L is large enough, then parents will always appear to follow a son-preferring fertility stopping rule, and they will invest in their boys such that they have at least one surviving son. If the disutility of sons dying is large enough relative to the inheritance functions, then parents could still have more than one surviving son. The household size effect is the same as above. As the proportion of boys increases, parents will want to decrease investment in sons to maximize the bequest per son. The sex composition effect is ambiguous for girls, depending on how much investment in boys drops as the proportion of boys rises. Proofs for the sex composition effect are shown in Appendix D. This model is then consistent with the empirical results on child mortality: stopping rules cause higher boy mortality in boy-first households via the sex composition effect and higher girl mortality in girl-first households via the household size effect.

One implication of the model is that we should only expect stopping rules and a child mortality gap for parents which own inheritable land. The RCH II does not ask about land ownership. However, the Indian National Family Health Survey (NFHS) does. I use the 2005/2006 round of the NFHS, a nationally representative survey of approximately 125,000 households, to compare households that do and do not own land. I run a similar estimation as in Equation (2) above.

(6)

$$\#ChildrenBorn_{ij} = \beta_{FB} FirstBornBoy + \beta_L OwnLand + \gamma FirstBornBoy * OwnLand + \beta_1 X_i + \beta_j State_j + e_{ij}$$

Here I add a variable for whether the household owns any land and an interaction term for owning land and having a first-born boy. X_i is the same as above, except that there is no variable for husband's education. The coefficient of interest is γ on FirstBornBoy*OwnLand. Table 6 shows that for women aged 35 and older a first-born boy in a landowning household predicts about 0.12 fewer children than a first-born boy in a non-landowning household. Thus, stopping rules appear to be more prevalent amongst households where inheritance matters.

I also find that the mortality gap is significantly larger in landowning households. In households that do not own land 5.5% of boys and 6.3% girls die between the ages of 1 and 60 months, yielding a mortality gap of about 14%. In households that own land 6.4% of boys and 7.9% of girls die between the ages of 1 and 60 months, yielding a mortality gap of about 25%. Thus, the child mortality gap in households that own land is almost double the gap of rural households that do not own land. These results indicate that inheritance could be a major factor in the child mortality gap via fertility decisions. However, the economic incentives from owning land may be due to reasons besides inheritance. For example, owning land increases the importance of sons remaining in the joint household and contributing to the family farm.

Table 6: OLS: Dependent Variable = Total Children Born⋆

First Boy*Own Land		-0.124**
		(0.036)
Own Land		0.163**
		(0.030)
First Boy	-0.335**	-0.284**
	(0.017)	(0.022)
Household Variables	Yes	Yes
State Dummies	Yes	Yes
N	35717	35717
Clusters	3814	3814
R^2	0.3156	0.3162

Significance levels : † : 10% ∗ : 5% ∗∗ : 1%

Notes: Errors clustered by sample unit in parentheses.
Data: Indian NFHS 2005/2006.
⋆Mother aged 35 and older.

5.3 Unintended Consequences of Granting Inheritance Rights to Women

In 2005 India amended the Hindu Succession Act in its constitution, giving girls a legal claim to inherit their parents' land. What this means is that if daughters are denied an equal share of their parents' land by their family, then they have the ability to go to court to claim this share. It is not that parents now have the legal option to leave a share of their land to their daughters, but that they are legally required to do so. I show that giving women inheritance rights can increase child mortality rates. First, I explain theoretically why this is so. Then I will use a difference-in-difference analysis to show that state reforms that gave women inheritance rights raised child mortality rates.

Using the model of inheritance, if women are granted equal inheritance rights, then $\alpha = 1$. The results below are for a small increase in α when α is small. That is, I will explain what happens when women who have few inheritance rights are given a marginal increase in their rights. I prove the following proposition in Appendix D:

Proposition 4. *For Z sufficiently large and α sufficiently small, $\frac{\partial k_B}{\partial \alpha} < 0$ and $\frac{\partial k_G}{\partial \alpha} < 0$. That is, an increase in the amount of land inherited by girls will lower parents' health investments in boys and girls, resulting in higher child mortality rates.*

If parents continue to only care about maximizing their bequests to their sons, then daughters' rights to inherit are bad for parents. Parents would prefer to give all their land to their sons, but have to give some to their daughters. In order to reduce the number of these burdensome daughters, parents will lower health expenditures for their daughters, resulting in higher child mortality. As long as parents have more than one boy, boys will also suffer if their sisters are given inheritance rights. The drop in inheritance per child will cause parents to decrease their investment per son, so that they can increase their bequest per son. See Appendix D for an analytical explanation. Intuitively, giving women inheritance rights is like a reduction in parents' inheritable land. They will want to split up this smaller amount of land amongst fewer sons compared to if they had a lot of land. Z must be sufficiently large, i.e. parents must gain a lot of utility from small increases in land per son. The drop in investment in daughters frees up income which will be devoted to sons unless the consequent loss of land per son outweighs the benefits of improving son mortality. I focus on slight increases in α where α is small because, as I will show below, women inherit little and the increase in inheritance rights probably only raised inheritance a small amount. The model predicts that even for small increases in inheritance rights, we should expect a rise in child mortality.

To understand the effects of inheritance rights on child mortality, I use the 1999 REDS data and changes in inheritance laws from a few states that granted women inheritance rights before 2005: Kerala in 1976, Andhra Pradesh in 1986, Tamil Nadu in 1989, and Maharashtra and Karnataka in 1994.[18] I call these states "reform states". Note that the reforms only applied to "Hindu" households, which includes Hindus, Jains, Sikhs, and Buddhists. In my discussion below, when I refer to Hindu households I implicitly include all four religions.

Just because a law is changed, does not mean that it will be enforced, as is evident by the existence of dowry in India even though it has been officially illegal since 1961. As noted in Reitmaier (2007), the state law changes had a small but positive effect on the probability that women actually do inherit land. Using the 1999 REDS data, I look at households where the male head of household became the head at his father's death. The survey tells us when the head became

[18] See Agarwal (1995) for a brief exposition on inheritance laws and women's rights over agricultural land. Kerala's reform was different from the other four states in that it removed the legal status of the joint family altogether giving parents and children equal legal shares in family property.

the head, whether he has any sisters, and whether these sisters inherited anything. Focusing on the reform states, and using a Pearson's chi-squared test I examine whether there is an increase in the proportion of women who inherit after the reform. The results are shown in Table 7. Restricting the sample to Hindu households where the father of the head of household owned land at his death, I find a statistically significant increase in the proportion of women who inherit from about 3.5% before the reforms to 9.1% afterwards. Legally, the reform only included Hindu women who married after the reform. If we restrict ourselves to women who married after the reform or women who were unmarried, we see a larger increase in proportion of women who inherit from about 1.4% before the reform to 13.5% afterwards. Even if one considers this a small increase in inheritance for women, parents may still have expectations that their young daughters will have a large probability of being able to inherit in the future. In any case, the model predicts increased child mortality for a small increase in inheritance.

Table 7: Proportion of "Hindu" Women who Inherit from Landowning Fathers Pre and Post Reform

	Pre-Reform	**Post-Reform**	**Difference**
All	0.035	0.091	0.056**
Observations	874	470	
Married After Reform or Unmarried	0.014	0.135	0.121*
Observations	70	126	

Significance levels : † : 10% ∗ : 5% ∗∗ : 1%

As initial evidence that the inheritance law reforms caused higher mortality rates, I present a simple triple-differences estimate. I focus on Karnataka and Maharashtra, which both changed their laws at the same time in 1994. I look at mean girl and boy mortality separately. The first difference is child mortality before and after 1994. The second difference is Karnataka and Maharashtra versus the rest of India minus the other reform states. The third difference is whether households are Hindus and own land or not. The results are show in Tables 8 and 9. We can see that while mortality rates fell for everyone after 1994, they only fell a very small amount for boys and girls in Karnataka and Maharashtra in Hindu landowning households. These results suggest that the granting inheritance rights to women raised mortality rates about 2%.

For more precise estimates of the impact of the state reforms, I estimate a series of difference-in-

Table 8: Triple Differences: Boy Mortality*100⋆. Karnataka and Maharashtra vs. Non-Reform States.

	Karnataka and Maharashtra	**Non-Reform States**	**Difference-in-Difference**
Hindu Landowning			
Pre 1994	4.33	8.39	
Post 1994	3.55	5.06	
Difference	0.78	3.34**	2.55**
	(0.72)	(0.36)	(0.81)
Non-Hindu and/or Non-Landowning			
Pre 1994	4.69	7.60	
Post 1994	2.34	4.76	
Difference	2.35**	2.84**	0.49
	(0.53)	(0.34)	(0.63)
			Triple Difference
			2.07*
			(1.02)

Significance levels : † : 10% ∗ : 5% ∗∗ : 1%

Notes: Errors clustered by household in parentheses. Sample weights used.

⋆ Child born at least 60 months before survey and died between the ages of 1 and 60 months.

Table 9: Triple Differences: Girl Mortality*100*. Karnataka and Maharashtra vs. Non-Reform States.

	Karnataka and Maharashtra	**Non-Reform States**	**Difference-in-Difference**
Hindu Landowning			
Pre 1994	5.19	10.51	
Post 1994	3.02	6.88	
Difference	2.18**	3.63**	1.45†
	(0.72)	(0.47)	(0.85)
Non-Hindu and/or Non-Landowning			
Pre 1994	4.64	8.88	
Post 1994	1.66	6.20	
Difference	2.98**	2.63**	-0.35
	(0.52)	(0.42)	(0.65)
			Triple Difference
			1.80†
			(1.07)

Significance levels : † : 10% ∗ : 5% ∗∗ : 1%

Notes: Errors clustered by household in parentheses. Sample weights used.

⋆ Child born at least 60 months before survey and died between the ages of 1 and 60 months.

difference equations. My first treatment group is children born after the reform in the reform states. The second treatment group restricts the post-reform children to Hindu households, those covered by the law changes within the reform states. The third treatment group restricts the post-reform children to land-owning households, those likely affected by inheritance law changes. Last, I focus on children born after the reforms in both Hindu and land-owning households. The estimation equations are as follows:

$$M_{S,ijt} = \gamma_S PostReform + \beta_1 X_i + \beta_j State_j + \beta_t YOB_t + e_{ijt} \tag{7}$$

$$M_{S,ijt} = \gamma_S PostReform * OwnLand + \beta_L OwnLand + \beta_1 X_i + \beta_j State_j + \beta_t YOB_t + e_{ijt} \tag{8}$$

$$M_{S,ijt} = \gamma_S PostReform * Hindu + \beta_H Hindu + \beta_1 X_i + \beta_j State_j + \beta_t YOB_t + e_{ijt} \tag{9}$$

(10)
$$M_{S,ijt} = \gamma_S PostReform*Hindu*OwnLand + \beta_L OwnLand + \beta_H Hindu + \beta_1 X_i + \beta_j State_j + \beta_t YOB_t + e_{ijt}$$

$M_{S,ijt}$ is the mortality outcome of child of sex S in household i, state j, born in year t. The Post Reform variable and terms interacted with Post Reform are my "treatment" variables. State ($State_j$) and year of birth (YOB_t) fixed effects are included to remove state differences in mortality rates as well as any time trend in mortality rates. Household and child characteristics, X_i, include mother age and education, caste, religion, and birth order. I also include standard of living dummies (low, medium, or high) to control for wealth differences. Note that I drop Kerala from my estimations because almost all of the births in my sample occur after Kerala's reform.

There are possible biases in the estimates. All of the reforms, except for Maharashtra, are in southern India, which has better child outcomes than the rest of India. Thus, we may expect lower mortality rates in general in these states. As noted in Roy (2008), women who are eligible to inherit via these state reforms have higher autonomy within their households. If women have a stronger desire to keep their children alive than men, we would also expect the reforms to improve child mortality. Although, according to Eswaran (2002)'s intrahousehold bargaining model of child mortality, we may expect the increased autonomy of mothers to improve son outcomes and worsen daughter outcomes. The intuition for this is that mothers are much more dependent on their sons for retirement security than are fathers. In any case, the bias is for improved son mortality, with possibly worse daughter mortality. That the reforms were passed in only some states may reflect that landowners, who would suffer from the reforms, are not as politically strong as in other states. This could indicate that landowners in reform states are not as wealthy as elsewhere and bias the results towards higher child mortality. I include controls based on the standard of living index as a proxy for wealth to help to account for some of this potential bias.

The results are shown in Table 10 and Table 11. γ_S is the coefficient of interest in each equation. We see that the coefficient is positive and statistically significant for both boy and girl mortality for several of the specifications. For the post-reform Hindu treatment group, in columns 3 and 4, the results are statistically significant for each of the specifications. This is the group that is specifically targeted by the reforms. The results remain a similar size when I focus on post-reform Hindu land-owning households in columns 7 and 8. However, they lose their statistical significance. The effects are of a similar magnitude as the effect of the sex outcome of the first pregnancy. The estimations show a rise in boy mortality of around 1% and a rise in girl mortality of around 0.5%. This is even with the potential bias from the reforms occurring in states in southern India or causing an improvement in women's autonomy. Thus, there is evidence that the legal reforms that had the intention of improving women's rights had the unintended consequence of raising child mortality rates.

Table 10: OLS: Dependent Variable = Boy or Girl Dead × 100*

	(1)		(2)		(3)		(4)	
	Boys	**Girls**	**Boys**	**Girls**	**Boys**	**Girls**	**Boys**	**Girls**
Post Reform*Hindu					1.223*	0.736*	0.908**	0.564†
					(0.312)	(0.330)	(0.314)	(0.332)
Hindu					-0.469*	-0.194	0.138	0.592*
					(0.211)	(0.230)	(0.220)	(0.240)
Post Reform	0.748*	0.359	0.931**	0.764*				
	(0.335)	(0.358)	(0.335)	(0.359)				
Religion Dummies	No	No	Yes	Yes				
Mother Age and Education	No	No	Yes	Yes	No	No	Yes	Yes
Caste Dummies	No	No	Yes	Yes	No	No	Yes	Yes
Wealth Dummies	No	No	Yes	Yes	No	No	Yes	Yes
Birth Order Dummies	No	No	Yes	Yes	No	No	Yes	Yes
State and Year Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
N	98556	92190	96611	90394	98556	92190	96756	90516

Significance levels : † : 10% ∗ : 5% ∗∗ : 1%

Notes: Errors clustered by household in parentheses.

⋆ Child born at least 60 months before survey and died between the ages of 1 and 60 months.

Table 11: OLS: Dependent Variable = Boy or Girl Dead × 100*

	(5)		(6)		(7)		(8)	
	Boys	**Girls**	**Boys**	**Girls**	**Boys**	**Girls**	**Boys**	**Girls**
Post Reform*Hindu*Own Land					0.907*	0.503	0.695	0.320
					(0.433)	(0.433)	(0.434)	(0.437)
Post Reform*Own Land	0.793†	0.401	0.676	0.346				
	(0.417)	(0.419)	(0.420)	(0.425)				
Hindu					-0.431*	-0.222	0.127	0.525*
					(0.203)	(0.220)	(0.212)	(0.231)
Own Land	0.542**	0.808**	0.584**	0.822**	0.571**	-0.817**	0.602**	0.819**
	(0.171)	(0.191)	(0.171)	(0.1935)	(0.170)	(0.190)	(0.172)	(0.191)
Religion Dummies	No	No	Yes	Yes				
Mother Age and Education	No	No	Yes	Yes	No	No	Yes	Yes
Caste Dummies	No	No	Yes	Yes	No	No	Yes	Yes
Wealth Dummies	No	No	Yes	Yes	No	No	Yes	Yes
Birth Order Dummies	No	No	Yes	Yes	No	No	Yes	Yes
State and Year Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
N	98556	92190	96756	90394	98556	92190	96756	90516

Significance levels : † : 10% * : 5% ** : 1%

Notes: Errors clustered by household in parentheses.

⋆ Child born at least 60 months before survey and died between the ages of 1 and 60 months.

6 Conclusion

This paper makes several contributions to the literature on fertility decisions and intrahousehold discrimination. It provides a theoretical framework to understand how economic incentives cause parents to follow fertility stopping rules and how these decisions will disadvantage girls on average. One new theoretical result is the effect of sibling sex composition on girl and boy mortality: boys want sisters and girls want brothers. I find that the sex of the first-born child explains about one quarter of the child mortality gap between boys and girls. This means that fertility stopping rules and the resulting resource discrimination are a major cause of excess girl mortality in India. I find that sex selective abortions should counteract the effects of stopping rules, lowering the child mortality gap. I also provide evidence to verify that health resources in the form of vaccinations are affected by stopping rules.

This paper makes some progress toward understanding the specific economic incentives driving fertility decisions and discrimination against girls. Dowries are likely to be a large burden for parents, and daughters would be less costly, and thus more likely to survive, if dowries did not exist. However, if women are a net economic burden to households, then the receipt of dowries by sons are unlikely to be of much benefit to parents. Thus, the economic advantage of sons is more likely due to their relatively high wages and contribution to the joint household throughout their lives. Another possibility is that parents' desire to maximize their bequest per son is responsible for fertility decisions that create the child mortality gap. Furthermore, if women are given inheritance rights, but parents prefer not to give their land to their daughters, then these new rights will have unintended consequences. In order to maintain large plots of land per son, parents reduce their health investments in their children, leading to higher boy and girl mortality.

Now that we know that girls suffer in India from the fertility decisions of their parents and girls' relatively poor future economic outlooks, what can we do about it? Although I have seen some efforts to reduce the burden of marriage costs,[19] it seems unlikely that dowries will be eradicated in India in the foreseeable future, even though the practice has been officially illegal for almost fifty

[19]SKDRDP in Dharmasthala, India, for example, has held several free mass weddings which they have made attractive by using the strong religious influence of the Dharmasthala temple.

years. Solutions will then need to focus on improving the economic situation of women or reducing the economic burdens of girls. The rise of women's microcredit groups is no doubt a start. Perhaps payments to households with girls tied to proof of medical care and education for girls is a viable solution, as in the Mexican Progresa program. As a start, this paper provides some insight into the mechanisms through which girls are disadvantaged in India.

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APPENDIX

A Proofs of Propositions 1, 2, and 3

Assuming parents have stopped their fertility at N children and taking equation (1) above and substituting in the budget constraints yields the following maximization problem.

$\max_{k_B,k_G} U_T = U_1(Y_1 - \pi N k_B - (1-\pi)N k_G - NF) + U_2(Y_2 + \pi N D p(k_B) - (1-\pi)NDp(k_G)) + A(p(k_B)\pi N + p(k_G)(1-\pi N))$

where k_B and k_G are health investments in boys and girls. $0 \leq \pi \leq 1$ is the proportion of boys. There are N total children. D is the size of the cost or benefit of daughters and sons respectively. Y_1 and Y_2 are parents' income in Periods 1 and 2. $0 \leq p(k_i) \leq 1$ is the number of surviving children of sex i, given health investment k_i. A is the weight on survival utility. U_i and p are positive and strictly concave functions.

To determine comparative statics with respect to π we must calculate the following first-order conditions:

First Order Condition 1$= \frac{\frac{\partial U_T}{\partial k_B}}{\pi N} = -U_1' + Dp'(k_B)U_2' + Ap'(k_B) = 0$

First Order Condition 2$= \frac{\frac{\partial U_T}{\partial k_G}}{(1-\pi)N} = -U_1' - Dp'(k_G)U_2' + Ap'(k_G) = 0$

Then, we must construct the matrix of partial derivatives:

$\frac{\partial^2 U_T}{\partial k_B^2} = \pi N U_1'' + \pi N D^2 p'(k_B)^2 U_2'' + Dp''(k_B)U_2' + Ap''(k_B) < 0$

$\frac{\partial^2 U_T}{\partial k_G \partial k_B} = \pi N U_1'' - \pi N D^2 p'(k_G)p'(k_B)U_2''$ is positive if D is large enough.

$\frac{\partial^2 U_T}{\partial k_B \partial k_G} = (1-\pi)NU_1'' - (1-\pi)ND^2 p'(k_G)p'(k_B)U_2''$ is positive if D is large enough.

$\frac{\partial^2 U_T}{\partial k_G^2} = (1-\pi)NU_1'' + (1-\pi)ND^2 p'(k_G)^2 U_2'' - Dp''(k_G)U_2' + Ap''(k_G)$, which can be positive or negative depending on whether:

$(1-\pi)ND^2 p'(k_G)^2 U_2'' - Dp''(k_G)U_2'$ is positive or negative. It is negative if D is large enough.

$\frac{\partial^2 U_T}{\partial k_B \partial \pi} = N(k_B - k_G)U_1'' + ND^2 p'(k_B)(p(k_B) + p(k_G))U_2'' < 0,$

since $k_B - k_G > 0$

$\frac{\partial^2 U_T}{\partial k_G \partial \pi} = N(k_B - k_G)U_1'' - ND^2 p'(k_G)(p(k_B) + p(k_G))U_2'' > 0$ if D is large enough.

Then

$$\frac{\partial k_B}{\partial \pi} = -\frac{Det\begin{vmatrix} \frac{\partial^2 U_T}{\partial k_B \partial \pi} & \frac{\partial^2 U_T}{\partial k_B \partial k_G} \\ \frac{\partial^2 U_T}{\partial k_G \partial \pi} & \frac{\partial^2 U_T}{\partial k_G^2} \end{vmatrix}}{Det\begin{vmatrix} \frac{\partial^2 U_T}{\partial k_B^2} & \frac{\partial^2 U_T}{\partial k_B \partial k_G} \\ \frac{\partial^2 U_T}{\partial k_G \partial k_B} & \frac{\partial^2 U_T}{\partial k_G^2} \end{vmatrix}} = -\frac{Det\begin{vmatrix} - & + \\ + & - \end{vmatrix}}{Det\begin{vmatrix} - & + \\ + & - \end{vmatrix}}$$

Multiplying out the numerator and simplifying, dividing by N, and letting $\pi N = B$ and $(1 - \pi)N = G$, we get:

$GD^2 U_1'' U_2''[(p'(k_B) + p'(k_G))(p(k_B) + p(k_G)) + (k_B - k_G)(p'(k_G)^2 + p'(k_G)p'(k_B))]$

$+ (A - DU_2')p''(k_G)[D^2 p'(k_B)(p(k_B) + p(k_G))U_2'' + (k_B - k_G)U_1'']$

Multiplying out the denominator and simplifying we get:

$BGD^2 U_1'' U_2''(p'(k_G)^2 + p'(k_B)^2 + 2p'(k_B)p'(k_G) + \frac{U_2' p''(k_G)}{BDU''2})$

$+ (A^2 - D^2 U_2'^2)p''(k_G)p''(k_B)$

$+ (A + DU_2')GD^2 U_2'' p'(k_G)^2 p''(k_B)$

$+ (A - DU_2')BD^2 U_2'' p'(k_B)^2 p''(k_G)$

$+ (A + DU_2')Gp''(k_B)U_1''$

$+ (A - DU_2')Bp''(k_G)U_1''$

Note that if $A > DU_2'$, that is, if the marginal utility of survival of an extra daughter is larger than the marginal consumption utility loss of an extra daughter in Period 2, then both the numerator and denominator are positive, making $\frac{\partial k_B}{\partial \pi} < 0$. By FOC2, this must be true: $\frac{U_1'}{p'(k_G)} = A - DU_2'$, which is positive because $\frac{U_1'}{p'(k_G)}$ is positive. And, thus, we have proved Proposition 1.

$$\frac{\partial k_G}{\partial \pi} = -\frac{Det\begin{vmatrix} \frac{\partial^2 U_T}{\partial k_B^2} & \frac{\partial^2 U_T}{\partial k_B \partial \pi} \\ \frac{\partial^2 U_T}{\partial k_G \partial k_B} & \frac{\partial^2 U_T}{\partial k_G \partial \pi} \end{vmatrix}}{Det\begin{vmatrix} \frac{\partial^2 U_T}{\partial k_B^2} & \frac{\partial^2 U_T}{\partial k_B \partial k_G} \\ \frac{\partial^2 U_T}{\partial k_G \partial k_B} & \frac{\partial^2 U_T}{\partial k_G^2} \end{vmatrix}} = -\frac{Det\begin{vmatrix} - & - \\ + & + \end{vmatrix}}{Det\begin{vmatrix} - & + \\ + & - \end{vmatrix}}$$

The denominator will be positive as above. Multiplying out the numerator and simplifying, dividing by N, and letting $\pi N = B$ and $(1 - \pi)N = G$, we get:

$BD^2 U_1'' U_2''[(k_B - k_G)(p'(k_G)p'(k_B) - p'(k_B)^2) - (p'(k_B) + p'(k_G))(p(k_B) + p(k_G))]$

$+ (A + DU_2')p''(k_B)[(k_B - k_G)U_1'' - D^2 p'(k_G)(p(k_B) + p(k_G))U_2'']$

It is not obvious whether the numerator is negative or positive. Note that if $k_B = k_G$, it must be negative. However, as k_B and k_G diverge, it is possible that the terms become positive. For D large enough, the second term will become negative and larger (in absolute value) than the first term, since it contains a D^3, while the first term only has a D^2. Thus, if D is large enough the numerator is negative and $\frac{\partial k_G}{\partial \pi} > 0$. This proves Proposition 2.

To understand what happens to incentives to continue having children, I alter the model slightly:

$U_T(k_B, k_G) = U_1(Y_1 - Bk_B - Gk_G - (B+G)F) + U_2(Y_2 + BDp(k_B) - GDp(k_G)) + A(p(k_B)B + p(k_G)G)$

So now there are explicitly B boys and G girls. What happens to incentives to have an extra child (assumed with 50% probability to be a boy and 50% probability of being a girl)?

The utility gain from having slightly more boys and slightly more girls is:

$\frac{\partial U_T}{\partial B} + \frac{\partial U_T}{\partial G} = (-k_B - k_G - F)U_1' + D(Bp(k_B) - Gp(k_G))U'2 + A(p(k_B) + p(k_G))$

If we raise the proportion of boys, as B goes up and G goes down, k_B goes down and k_G goes up (from above). Although it is ambiguous what happens to Period 1 and survival marginal utility (depending on how much k_B decreases and how much k_G increases), Period 2 marginal utility must fall. As long as D is large enough, this effect will dominate, causing parents to gain less utility from an extra child. Thus, we have proved Proposition 3.

B Savings and Credit

To illustrate as simply as possible how allowing parents to borrow against future dowry payments may reverse Proposition 1, I simplify the model by focusing solely on boys, so that the maximization problem becomes:

$\max_{k_B, S} U_T(k_B, S) = U_1(Y_1 - Bk_B - S) + U_2(Y_2 + BDp(k_B) + RS) + ABp(k_B)$

where R is the rate of interest + 1.

The first-order conditions are:

$FOC1 = \frac{\partial U_T}{\partial k_B} = -BU_1' + DBp'(k_B)U_2' + ABp'(k_B) = 0$

$FOC2 = \frac{\partial U_T}{\partial S} = -U_1' + RU_2' = 0$

Note that parents will always set S to satisfy

$R = \frac{U_1'}{U_2'} = Dp'(k_B) + \frac{Ap'(k_B)}{U_2'}$.

Parents invest in their sons until the return from investing in sons is equal the return from saving. If $A = 0$, i.e. parents only care about the economic benefits of sons, then for however many sons they have, they will invest in their sons up until $Dp'(k_B) = R$. Thus, regardless of the number of sons, parents will not change their investment and child mortality will not change. If $A > 0$, this is no longer an equilibrium. To see why, note that if B increases, ceteris paribus, U_1' goes up and U_2' goes down. If S is set such that again $R = \frac{U_1'}{U_2'}$, via borrowing against future child benefits, then $Dp'(k_B) + \frac{Ap'(k_B)}{U_2'} > R$, since U_2' is smaller than before. Thus, in equilibrium, k_B must rise somewhat, giving the exact opposite result as in Proposition 1. Of course, if R is sufficiently large, parents will never borrow against future child benefits, and Proposition 1 will again hold. Since Proposition 2 stems from the overall wealth increase of extra sons, and not from the intra-period resource re-allocation as for sons, the introduction of savings and credit should not change Proposition 2.

C Exogeneity of the Sex Outcome of the First Pregnancy

As noted by, for example, Ebenstein (2007), the first pregnancy in India has a normal male/female sex ratio (i.e. in the range of 1.04-1.07 males per female)[20], while later births show an increase in the sex ratio, indicating the use of selective abortion. As Ebenstein (2007) points out, we also see the phenomenon of increasing sex ratios among higher birth orders in China, where the first birth has an approximately normal sex ratio (also see Das Gupta (2005)). We cannot treat later births as random because some households choose to have these children only if the fetus is male. If parents do selectively abort their first pregnancy this will present a selection bias for the reduced form estimates in the paper. For example, if parents who selectively abort are those that are richer or better educated or have better access to health facilities, then we would expect that in families with first-born girls, girls are more likely to die for reasons completely outside of resource discrimination within a household. Figure 2 is a state-by-state graph of the sex ratio of children aged 0-6 from

[20] See Chahnazarian (1988) for a review of literature on the biologically normal sex ratio at birth and Parazzini et al. (1998) on recent global trends in the sex ratio at birth.

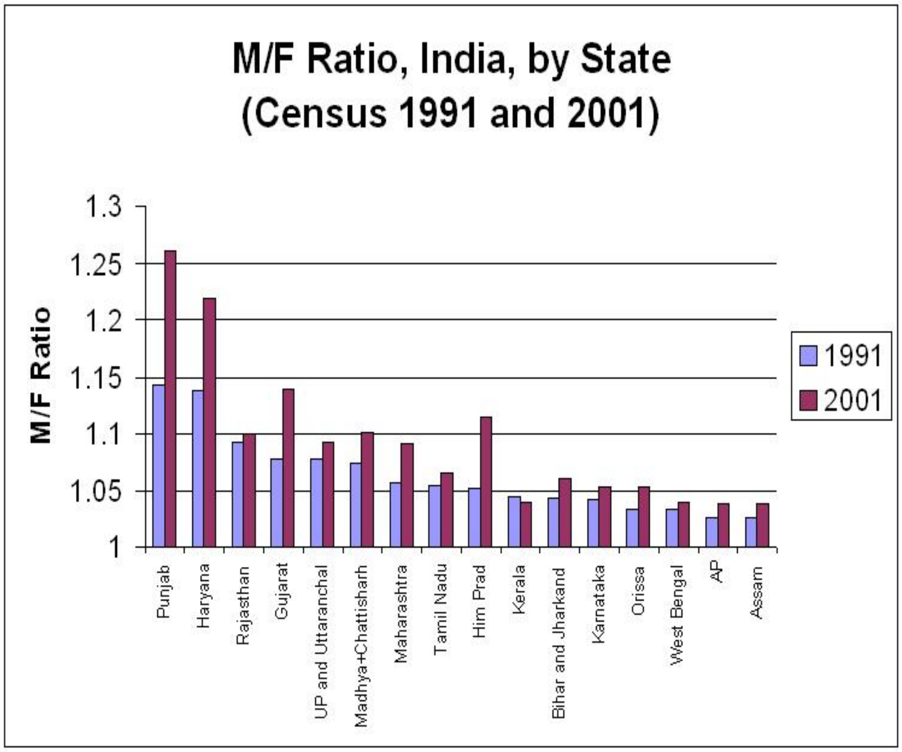

Figure 2: Male/Female Ratio by state, 1991 and 2001, Age 0-6

the 1991 and 2001 Indian censuses. The sex ratio is high for many of the states, so the presence of selective abortion is a concern.

Table 12 presents estimates of the sex ratio at birth by birth order using the RCH II. The sex ratio for first-borns amongst all women surveyed is 1.089, a bit above what we would expect to occur naturally. This number is deceptive because older interviewed women potentially have recall bias about their birth history. Recall bias occurs when parents who had a first-born daughter, but the daughter died during infancy, do not remember or report the first-born daughter. There is also survival bias, in the sense that having a girl first increases the total number of children born, which increases the chance of the mother dying. Having children will also increase morbidity and indirectly lead to higher mortality rates (Jejeebhoy and Rao, 1995). If, instead, we look at more recent births, where mother mortality is lower and recall and survival bias is lessened, then the sex ratio at birth for first-borns who were 0-10 years old at the time the survey was taken, falls to 1.066, in the range considered normal, while remaining high for higher birth orders.

Table 12: RCH II: M/F Ratio by Birth Order

Birth Order	All Ages	95% CI	Born Within Ten Years of Survey	95% CI
1st	1.089	(1.083, 1.095)	1.066	(1.056, 1.076)
2nd	1.088	(1.081, 1.095)	1.091	(1.080, 1.102)
3rd	1.100	(1.091, 1.109)	1.094	(1.081, 1.107)
4th	1.091	(1.080, 1.102)	1.091	(1.075, 1.108)
5th	1.070	(1.052, 1.082)	1.073	(1.054, 1.094)

Notes: sample weights used, no twins

Figure 3 illustrates what happens to the sex ratio at birth over time in the RCH II. Sex ratios at birth are broken down into 5-year birth groups, separated by birth order.[21] For example, "86-82" on the x-axis in the graph includes all children born in 1982 through 1986. The sex ratio for first-borns remains normal for more recent births, and then rises substantially going back to births in the 1980s. Selective abortion did not become widely available in India until the 1990s, making it unlikely that selective abortion is the cause of this rise in the sex ratio. Second and third order births have higher than normal sex ratios going back to the early 1990s, approximately the time sex selective abortion was becoming available, and approximately normal sex ratios for a brief period before this. The sex ratio is much higher in the more distant past. The sex ratio is always lower for 2nd and 3rd order births in the more distant past (before 1990) than 1st order births, and 4th and higher order births are even lower before 1990. This fits the survival bias story: we expect a first-born girl to have a stronger impact on fertility than a 2nd or 3rd-born girl, while 4th and higher order female births should see little or no impact on fertility. Thus, if a rise in fertility causes a rise in mortality risk, we expect survival bias to appear more strongly for lower birth-orders. This pattern also fits recall bias: it is more likely for a mother to misreport births that are more distant in her past, so that we may expect more recall bias for first pregnancies than later pregnancies.[22]

Another way to check whether survival or recall bias is happening is to see if the sex ratio at birth in surveys is rising in the 1980s as surveys are taken later and later from the 1980s. As the cumulative effects of extra births on health kill more and more women or more women

[21] The 95% confidence interval is approximately +/- 0.01, where it is slightly smaller for more recent births and larger for more distant births. The first data point includes all births in 1985 and before.

[22] A similar pattern of rising sex ratios for births more distantly in the past has been reported in Bangladesh (Majumder et al., 1997).

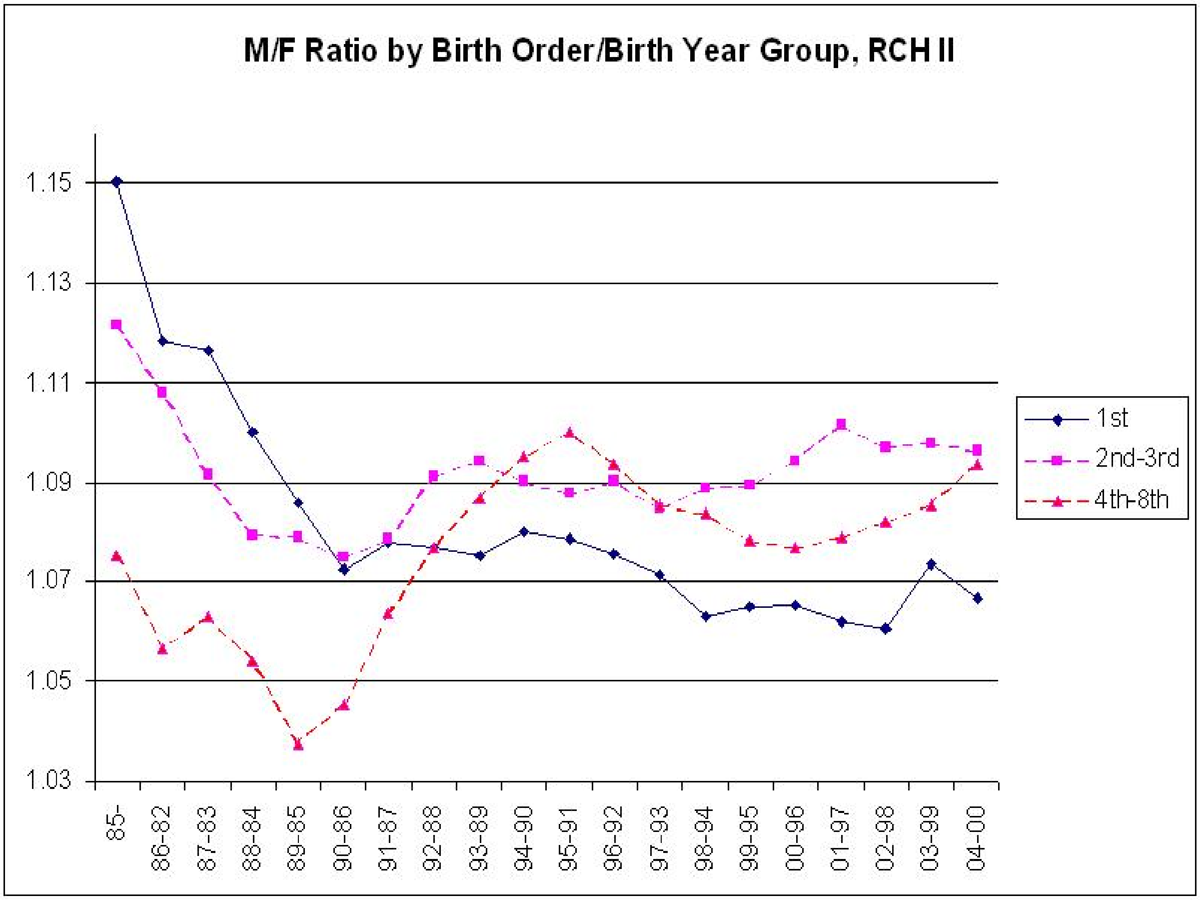

Figure 3: Male/Female Ratio by Birth Order/5-Year Birth Group, RCH II, All India

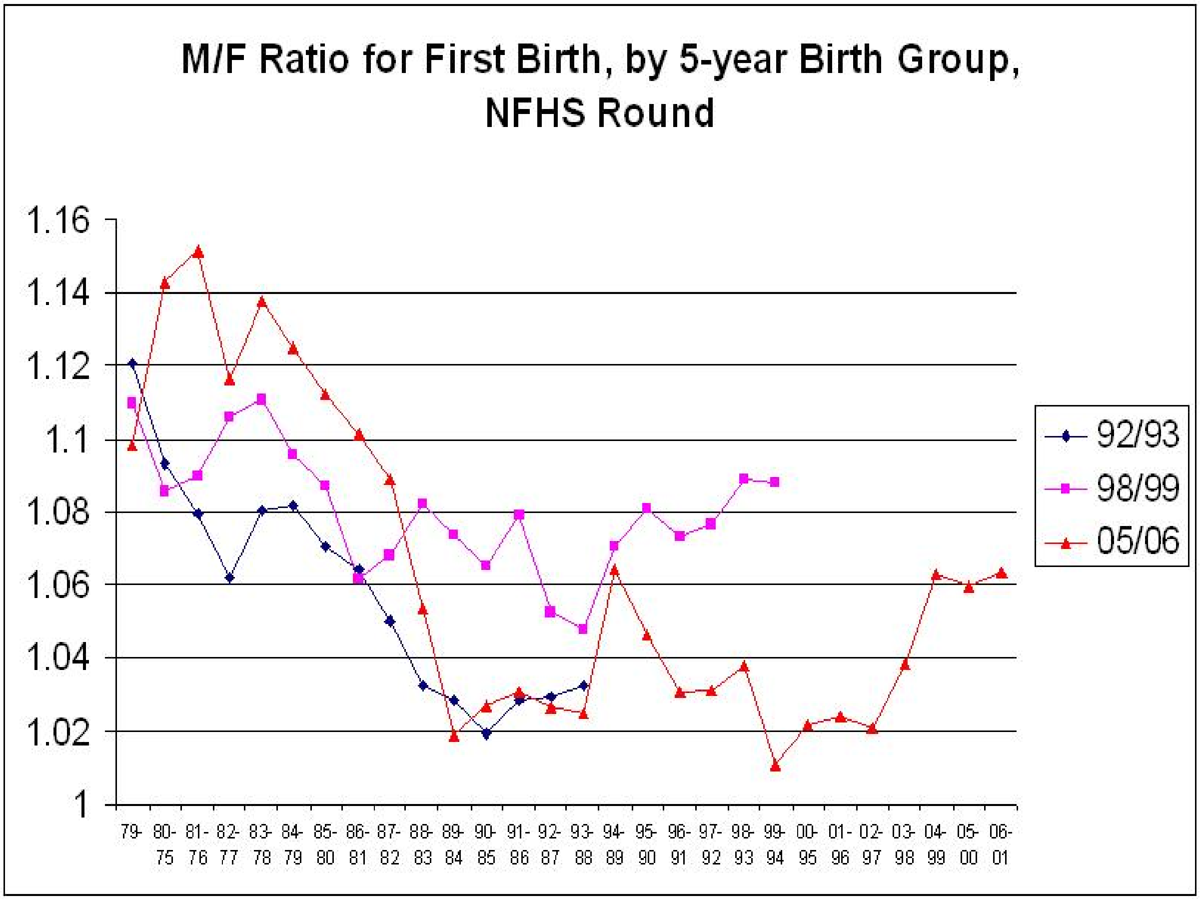

Figure 4: Male/Female Ratio for First Birth by 5-Year Birth Group/NFHS Round, All India

forget or choose to ignore their true first births, we should see such a rise. Figure 4 uses the three Indian National Family Health Survey rounds in 1992/1993, 1998/1999, and 2005/2006, all of approximately 80,000 ever-married women aged 13-50, to look at trends in the sex ratios of first-borns over time. I continue to use 5-year birth groups.[23] The relatively fewer number of births in the NFHS compared the RCH II cause wider fluctuations and larger confidence intervals. There is a trend of a rising sex ratio at birth in the 1980's, the more recent the survey. This trend is consistent with survival and recall bias.

By using some back of the envelope calculations, I estimate how large recall and survival bias must be to account for the difference in the RCH II between 1.089 and the 1.07 birth ratio which

[23]The 95% confidence interval is approximately +/- 0.025, where it is slightly smaller for more recent births and larger for more distant births. The first data point includes all births in 1979 and before.

would be considered "normal". The population of women who had a daughter first is estimated to be 1.7% smaller than it should be (i.e. than if they instead had a son first). If we assume that 0.5% of women die from childbirth, and mother's with a first-born daughter have 1/3 of an extra child, then this will cause only about 1/6 of a percent of women with a first-born girl to die. Thus, death during childbirth (survival bias) can only account for about 10% of the fewer women who have a daughter first. The rest must be due to recall bias or increases in post-birth mortality.

The bottom line is that, overall in India, it does not appear that selective abortion is occurring amongst first-borns, and is at least not the cause for the seemingly high sex ratio of first-borns in the cross-sectional data set. Children born to households where the mother dies are worse off, and likely to have higher child mortality rates than if their mother had survived. If survival bias is present, we expect this to bias the estimates such that we underestimate the decrease in girl mortality of having a first-born boy (raising γ_{girl}) and overestimate the increase in boy mortality of having a first-born boy (raising γ_{boy}). If recall bias is instead present because parents do not report first-born girls who died young, again we will be underestimating the decrease in girl mortality from having a first-born son (raising γ_{girl}), while misrecorded first-born sons, assuming they are from a family with high mortality, will cause an overestimate in the rise in boy mortality from having a first-born son (raising γ_{boy}). Both biases work in the same direction.

India has large differences in sex ratios across states. For example Punjab State has the worst child (age 0-6) sex ratio in India and the 1991 Indian census estimated this ratio at 1.14, rising to 1.26 in 2001. Thus, it is important to calculate the sex ratio at birth by state to make sure that some states with low sex ratios (e.g. Kerala) are not masking the sex ratios of states like Punjab. Table 13 presents sex ratios for the larger states of India for births within ten years of being surveyed for the RCH II. The small states are not shown because their low sample size and correspondingly large confidence intervals make them uninformative. About half of the states have a sex ratio of first-borns above 1.07 (although 1.07 is within most of the states' confidence intervals). In order to ensure that the estimates in the paper are robust to the possibility that sex selective abortion is occurring amongst first-borns in the states with first-born sex ratios above 1.07, the regressions are estimated with just the states with sex ratios below 1.07 in Table 13 and the results are similar

Table 13: M/F Ratio by "Large" Indian State, Age 0-10 at time of survey

State	1st Born	95% CI
Jammu & Kashmir	1.364	(1.280, 1.454)
Uttaranchal	1.109	(1.041, 1.181)
Chhattisgarh	1.108	(1.048, 1.172)
Karnataka	1.108	(1.064, 1.154)
Assam	1.105	(1.055, 1.158)
Rajasthan	1.101	(1.064, 1.140)
Haryana	1.100	(1.052, 1.151)
Himachal Pradesh	1.079	(1.010, 1.120)
Kerala	1.076	(1.016, 1.139)
West Bengal	1.071	(1.020, 1.124)
Uttar Pradesh	1.064	(1.038, 1.090)
Punjab	1.064	(1.012, 1.118)
Madhya Pradesh	1.061	(1.028, 1.096)
Tamil Nadu	1.061	(1.021, 1.102)
Andhra Pradesh	1.046	(0.998, 1.096)
Bihar	1.040	(1.007, 1.075)
Maharashtra	1.039	(1.001, 1.079)
Gujarat	1.031	(0.987, 1.076)
Arunachal Pradesh	1.029	(0.974, 1.088)
Orissa	1.012	(0.973, 1.052)

Notes: RCH II, sample weights used, no households with first-born twins, "large" states are those with more than 8500 respondents

to those reported in Table 3. The results are shown in Table 14.

Table 14: OLS: Dependent Variable = Percent of Boys or Girls Dead × 100 ⋆

States with sex ratio of first borns < 1.07 Mother Age ≥ 35, Birth Order ≥ 2		
Variable	**Boys**	**Girls**
FirstBornBoy	0.585**	-0.298†
	(0.133)	(0.158)
Age Mother	0.046*	0.138**
	(0.022)	(0.028)
Education Mother	-0.143**	-0.239**
	(0.016)	(0.023)
Education Father	-0.191**	-0.310**
	(0.016)	(0.022)
Rural	0.567**	1.343**
	(0.146)	(0.201)
Constant	1.645**	-0.166
	(2.881)	(1.483)
Religion Dummies	Yes	Yes
Caste Dummies	Yes	Yes
State Dummies	Yes	Yes
N	64283	56878
Clusters	490	490
R^2	0.023	0.035

Significance levels : † : 10% ∗ : 5% ∗∗ : 1%

Notes: Errors clustered by primary sampling units in parentheses, no first-born twins households.

⋆ Percentage of children born at least 60 months before survey and died between the ages of 1 and 60 months.

D Proofs for Model of Inheritance: Linear Preferences Over Inheritance

$$U_T = \begin{cases} U_1(c) + Z(\frac{L}{p(k_B)\pi N + \alpha p(k_G)(1-\pi)N}) + A(p(k_B)\pi N + p(k_G)(1-\pi)N) & \text{if } p(k_B)\pi N \geq 1 \\ \\ U_1(c) - J(L) + A(p(k_B)\pi N + p(k_G)(1-\pi)N) & \text{if } p(k_B)\pi N < 1 \end{cases}$$

subject to the following budget constraint: $c \leq Y - \pi N k_B - (1-\pi)N k_G - NF$

If J is sufficiently large, then parents will continue to have children until they have at least one surviving son. Of course, if parents continue having girls, eventually their resources will run out so that even if they had a boy they could not keep him alive, so that there is some maximum number of girls parents will have until they must stop having children. Thus, the model predicts son-preferring stopping rule behavior.

I prove the following below:

- 1. If Z is sufficiently large and $\alpha = 0$ and parents have more than one surviving boy, then $\frac{\partial k_B}{\partial \pi} < 0$.
- 2. If Z is sufficiently largeand $\alpha = 0$, and parents have more than one surviving boy, then $\frac{\partial k_G}{\partial \pi}$ is ambiguous.
- 3. Granting inheritance rights to girls will cause parents to lower their investment in girls and boys.

Below, I assume parents have followed a son-preferring stopping rule, have at least one surviving son and $\alpha = 0$. For ease of notation I will occasionally let B = πN and G = $(1-\pi)N$.

To determine comparative statics with respect to π we must calculate the following first-order conditions:

First Order Condition 1$= \frac{\frac{\partial U_T}{\partial k_B}}{B} = -U_1' - \frac{ILp'(k_B)}{(Bp(k_B))^2} + Ap'(k_B) = 0$

First Order Condition 2$= \frac{\frac{\partial U_T}{\partial k_G}}{G} = -U_1' + Ap'(k_G) = 0$

Then, we must construct the matrix of partial derivatives:

$\frac{\partial^2 U_T}{\partial k_B^2} = BU_1'' - \frac{IL}{B^2}\frac{p''(k_B)p(k_B)^2 - 2p'(k_B)^2 p(k_B)}{p(k_B)^4} + Ap''(k_B) > 0$ if Z is sufficiently large.

$\frac{\partial^2 U_T}{\partial k_G \partial k_B} = \frac{\partial^2 U_T}{\partial k_B \partial k_G} = GBU_1'' < 0.$

$\frac{\partial^2 U_T}{\partial k_G^2} = GU_1'' + Ap''(k_G) < 0.$

$\frac{\partial^2 U_T}{\partial k_B \partial \pi} = N(k_B - k_G)U_1'' + 2\frac{ILp'(k_B)}{(Np(k_B))^2}\frac{1}{\pi^3} > 0$, if Z is sufficiently large.

$\frac{\partial^2 U_T}{\partial k_G \partial \pi} = N(k_B - k_G)U_1'' < 0$ if $k_B - k_G > 0$.

Then

$$\frac{\partial k_B}{\partial \pi} = -\frac{Det\begin{vmatrix} \frac{\partial^2 U_T}{\partial k_B \partial \pi} & \frac{\partial^2 U_T}{\partial k_B \partial k_G} \\ \frac{\partial^2 U_T}{\partial k_G \partial \pi} & \frac{\partial^2 U_T}{\partial k_G^2} \end{vmatrix}}{Det\begin{vmatrix} \frac{\partial^2 U_T}{\partial k_B^2} & \frac{\partial^2 U_T}{\partial k_B \partial k_G} \\ \frac{\partial^2 U_T}{\partial k_G \partial k_B} & \frac{\partial^2 U_T}{\partial k_G^2} \end{vmatrix}} = -\frac{Det\begin{vmatrix} + & - \\ - & - \end{vmatrix}}{Det\begin{vmatrix} + & - \\ - & - \end{vmatrix}}$$

Thus, it must be the case that $\frac{\partial k_B}{\partial \pi} < 0$. This proves that if Z is sufficiently large, and parents have more than one surviving boy, then $\frac{\partial k_B}{\partial \pi} < 0$

For girls,

$$\frac{\partial k_G}{\partial \pi} = -\frac{Det\begin{vmatrix} \frac{\partial^2 U_T}{\partial k_B^2} & \frac{\partial^2 U_T}{\partial k_B \partial \pi} \\ \frac{\partial^2 U_T}{\partial k_G \partial k_B} & \frac{\partial^2 U_T}{\partial k_G \partial \pi} \end{vmatrix}}{Det\begin{vmatrix} \frac{\partial^2 U_T}{\partial k_B^2} & \frac{\partial^2 U_T}{\partial k_B \partial k_G} \\ \frac{\partial^2 U_T}{\partial k_G \partial k_B} & \frac{\partial^2 U_T}{\partial k_G^2} \end{vmatrix}} = -\frac{Det\begin{vmatrix} + & + \\ - & - \end{vmatrix}}{Det\begin{vmatrix} + & - \\ - & - \end{vmatrix}}$$

The denominator will be negative as above. The numerator could be positive or negative. Thus, the results are ambiguous for girls.

What happens if α rises? Parents would continue to follow stopping rule behavior as above. However, for simplicity, assume that the reform takes place after parents have made their fertility decisions and before they have made their health investment decisions. $\alpha = 0$ in the case above where women have no inheritance rights. In particular, the analysis is much simpler for a rise in α when α is small.

Thus, I will prove: If Z is large enough and α is small enough, then granting inheritance rights to girls (raising α) will cause parents to lower their investment in girls and boys: $\frac{\partial k_B}{\partial \alpha} < 0$ and $\frac{\partial k_G}{\partial \alpha} < 0$.

To determine the comparative statics, we need to revisit the first order conditions and partial derivatives above.

First Order Condition 1= $\frac{\frac{\partial U_T}{\partial k_B}}{B} = -U_1' - \frac{ILp'(k_B)}{(Bp(k_B)+\alpha Gp(k_G))^2} + Ap'(k_B) = 0$

First Order Condition 2= $\frac{\frac{\partial U_T}{\partial k_G}}{G} = -U_1' - \frac{\alpha ILp'(k_G)}{(Bp(k_B)+\alpha Gp(k_G))^2} + Ap'(k_G) = 0$

Then, we must construct the matrix of partial derivatives:

$\frac{\partial^2 U_T}{\partial k_B^2} = BU_1'' - IL\frac{p''(k_B)(Bp(k_B)+\alpha Gp(k_G))^2 - 2Bp'(k_B)^2(Bp(k_B)+\alpha Gp(k_G))^2}{(Bp(k_B)+\alpha Gp(k_G))^4} + Ap''(k_B) > 0$ if Z is sufficiently large.

$\frac{\partial^2 U_T}{\partial k_G \partial k_B} = \frac{\partial^2 U_T}{\partial k_B \partial k_G} = GBU_1'' + \frac{\alpha BGILp'(k_B)p'(k_G)}{2(Bp(k_B)+\alpha Gp(k_G))^3} < 0$ if α is small.

$\frac{\partial^2 U_T}{\partial k_G^2} = GU_1'' - \alpha IL\frac{p''(k_G)(Bp(k_B)+\alpha Gp(k_G))^2 - 2\alpha Gp'(k_G)^2(Bp(k_B)+\alpha Gp(k_G))^2}{(Bp(k_B)+\alpha Gp(k_G))^4} + Ap''(k_G) < 0$ if α is small.

$\frac{\partial^2 U_T}{\partial k_B \partial \alpha} = \frac{2BILp'(k_B)Gp(k_G)}{(Bp(k_B)+\alpha Gp(k_G))^3} > 0$

$\frac{\partial^2 U_T}{\partial k_G \partial \alpha} = -ILp'(k_G)\frac{(Bp(k_B)+\alpha Gp(k_G))^2 - 2\alpha Gp(k_G)(Bp(k_B)+\alpha Gp(k_G))}{(Bp(k_B)+\alpha Gp(k_G))^4} < 0$ if α is small enough or if the number of surviving girls is less than the number of surviving boys.

For boys,

$$\frac{\partial k_B}{\partial \alpha} = -\frac{Det\begin{vmatrix} \frac{\partial^2 U_T}{\partial k_B \partial \alpha} & \frac{\partial^2 U_T}{\partial k_B \partial k_G} \\ \frac{\partial^2 U_T}{\partial k_G \partial \alpha} & \frac{\partial^2 U_T}{\partial k_G^2} \end{vmatrix}}{Det\begin{vmatrix} \frac{\partial^2 U_T}{\partial k_B^2} & \frac{\partial^2 U_T}{\partial k_B \partial k_G} \\ \frac{\partial^2 U_T}{\partial k_G \partial k_B} & \frac{\partial^2 U_T}{\partial k_G^2} \end{vmatrix}} = -\frac{Det\begin{vmatrix} + & - \\ - & - \end{vmatrix}}{Det\begin{vmatrix} + & - \\ - & - \end{vmatrix}}$$

Both the denominator and numerator must be negative. Thus, $\frac{\partial k_B}{\partial \alpha} < 0$.

For girls,

$$\frac{\partial k_G}{\partial \alpha} = -\frac{Det\begin{vmatrix} \frac{\partial^2 U_T}{\partial k_B^2} & \frac{\partial^2 U_T}{\partial k_B \partial \alpha} \\ \frac{\partial^2 U_T}{\partial k_G \partial k_B} & \frac{\partial^2 U_T}{\partial k_G \partial \alpha} \end{vmatrix}}{Det\begin{vmatrix} \frac{\partial^2 U_T}{\partial k_B^2} & \frac{\partial^2 U_T}{\partial k_B \partial k_G} \\ \frac{\partial^2 U_T}{\partial k_G \partial k_B} & \frac{\partial^2 U_T}{\partial k_G^2} \end{vmatrix}} = -\frac{Det\begin{vmatrix} + & + \\ - & - \end{vmatrix}}{Det\begin{vmatrix} + & - \\ - & - \end{vmatrix}}$$

The denominator is negative. The numerator is negative if $\frac{\partial^2 U_T}{\partial k_B^2} * \frac{\partial^2 U_T}{\partial k_G \partial \alpha} < -\frac{\partial^2 U_T}{\partial k_B \partial \alpha} * \frac{\partial^2 U_T}{\partial k_G \partial k_B}$. This holds true if Z is large enough and α is small enough.