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Abstract

This paper argues that the social institutions of lineage maintenance, patrilocality and joint families have a significant role in explaining sex differences in survival and health outcomes in rural India, even when parents do not treat boys and girls differently. Tests using panel data from rural households confirm this explanation, which accounts for 7 percent of excess female mortality in Haryana and Rajasthan and 4 percent in Punjab. An institutional explanation suggests limits on the role for public policy in addressing large sex differences in health and mortality outcomes.

Keywords: Strategic bequests. Joint family. Fertility choice. Gender discrimination. Sex ratio.
JEL Codes: H31, J12, J13, J16, O15.

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1 Introduction

Boys and girls in India experience large differences in survival and health outcomes. The 2001 Census reports that the sex ratio for children under six years of age is 927 girls per thousand boys, one of the largest differences in survival outcomes in the world. Among surviving children, boys are more likely than girls to receive immunizations, medical attention and adequate nutrition (Pande 2003). An extensive literature has addressed these persistent gender differences, identifying various motivations such as differential returns in the labor market (Rosenzweig and Schultz 1982) and asymmetrical preferences due to culture or tradition (Sen 1990). These differences cause a sex bias both in labor market participation as well as in parents' investments in their children's health and education (Sen and Sengupta 1983). However, these explanations do not account fully for all aspects of health and mortality differences. In this paper, I show that bequests of land and associated fertility behavior in an agricultural society can be a significant driver of differential health and survival outcomes for boys and girls, even if parents do not treat daughters and sons differently.

Much of the existing literature suggests that parents actively discriminate in favor of boys through sex-selective foeticide and infanticide, as well as differences in provision of food and healthcare.[1] However, the evidence suggests that these explanations are incomplete because the estimated number of excess female deaths due to foeticide or infanticide do not account for the observed sex ratio (Dreze and Sen 2002). Despite arguments that parents actively discriminate against daughters in allocating nutrition and health resource, tests of intrahousehold allocation using recent data fail to reveal significant bias in behavior (Griffiths, Matthews, and Hinde 2002).

[1] See Das Gupta (1987) and the extensive literature cited in Miller (1981). The specific behaviors influencing gender differences in survival and health outcomes include abortion if pre-natal diagnostic testing reveals the foetus is female, infanticide if the newborn is a girl, and discrimination in the allocation of food and medical care in favor of boys throughout infancy and childhood. Policy responses have therefore sought to directly address these actions. In 1994, the Pre-Natal Diagnostic Techniques Act regulated the use of ultrasound machines and banned the use of "techniques for the purpose of pre-natal sex determination leading to female foeticide". State governments in Delhi and Haryana launched the "Ladli" scheme offering payments to low-income parents whose daughters survive childhood and achieve certain educational targets. Under the "Palna" scheme, the central government established "Cradle Baby Reception Centres" in each district where parents could leave unwanted girls for either future adoption or rearing in state-run orphanages.

Instead, Basu (1989) and Arnold, Choe, and Roy (1998) present evidence that son-preference manifests itself predominantly in fertility behavior so that the resulting family structure is unfavorable to girls. Clark (2000) and Jensen (2003) argue that this fertility behavior takes the form of "stopping rules" where parents have children until a certain number of boys are born. Under such rules, the average girl in the population will have systematically more siblings than the average boy, leading to lower resource allocation and poorer outcomes even with equitable parent behavior in distribution. Rosenblum (2010) and Barcellos, Carvalho, and Lleras-Muney (2010) estimate that stopping rules have a significant impact on differential outcomes for girls compared to boys. The first contribution of this paper is to propose a plausible explanation for the origin of these stopping rules.

Another line of evidence suggests that sex imbalances are not uniform across all households. Rosenzweig and Schultz (1982) argue that discrimination against girls is driven by asymmetry in the economic or social marketplace, which would suggest that the worst outcomes should be observed in the most destitute families where the marginal value of an additional son is greatest. However, Census and National Sample Survey data shows the sex ratio is worse in Indian states such as Punjab and Haryana where land forms a large part of family assets (Figure 1) and where income from agriculture is high (Figure 2). Mahajan and Tarozzi (2007) report that gender differences in nutrition and health outcomes increased in the 1990s, a period of rapid economic growth. Das Gupta (1987) as well as Chakraborty and Kim (2009) find that the difference between girls and boys is greater in middle class and higher caste households compared to lower class and lower caste households. If economic considerations drive discriminatory behavior, why are outcomes for girls relatively worse in agriculturally productive regions and among comparatively prosperous households? Addressing these contradictory findings is the second contribution of this paper.

Household level data investigated in this paper indicates that girls experience worse outcomes in large, multi-generational families known as "joint" families, which are common in rural farming communities. Caldwell's (1984) framework sheds light on various family structures in India. A "nuclear family" is formed when a couple leaves their parents' home upon marriage to form a household with their unmarried, typically minor, children. In a “stem family”, two married couples cohabit in a household together. The younger husband is the son of the older couple. Finally, a “joint-stem family” refers to a family where an older patriarch and his wife live with two or more adult sons, along with their wives and minor children.[2] In the Rural Economic and Demographic Survey (REDS 1999), the child sex ratio was 0.816 girls per boy in joint families, compared to 0.912 girls otherwise. Why this would be is not clear since recent research has shown that children in joint families benefit from higher levels of public good provision (Edlund and Rahman 2005). Proposing an explanation that is consistent with worse outcomes for girls in joint families is the third contribution of this paper.

I construct a model of bequest and fertility behavior among rural, land-owning families in a patrilocal society. In most regions of India, adult daughters leave their natal family at the time of marriage for their husband’s home, and are considered members only of their family of marriage. Consequently, they have few inheritance rights in their parents’ family, either in law or in practice, since any land given to them would be lost to the family lineage (Agarwal 1998; Chen 2000; Botticini and Siow 2003). The joint family head divides the land bequest among the remaining claimants, who are his adult sons. In so doing, the head is motivated by a desire to retain land within the family line carried through by his male descendants.

Why is land preservation so important in an agricultural society, particularly compared to more liquid assets such as cash, or those that are more directly consumed such as livestock? First, land is a fixed, immovable asset that cannot be lost or stolen. Thus, unlike wage employment, land offers a source of permanent income either through sale or direct consumption of the product. This has important consequences in a society with little formal social insurance. For example, Rose (1999) reports that controlling for size of asset holdings, child survival outcomes are significantly better in land owning families. Second, farmers who cultivate their own land do not face classic agency problems and are motivated to exert maximum effort into production (Banerjee, Gertler,

[2] In this paper, I use “independent family” instead of nuclear family, and “joint family” as a shorthand for a joint-stem family. Additionally, I differentiate between a “family” and “household” in the data, so co-residence within the household is not a requirement for membership of a joint or stem family.

and Ghatak 2002). Third, the advantages of land compared to other types of assets are recognized by other agents in the village economy. For example, Feder and Onchan (1987) show that land ownership improves access to credit, even if it not directly linked to farm investments. These reasons suggest that well-being of the lineage is symbiotic with preservation of land. Indeed, in a pioneering study of Indian villages, Srinivas (1976) wrote

> A man was acquiring land not only for himself but for his descendants... while a man may have had his descendants in mind when buying land he also knew it would be divided after his death... but even worse than division of land among descendants was not having any. That meant the end of the lineage, a disaster which no one liked even to contemplate.

Land possession, control and preservation is a significant factor influencing behavior within rural families. With land sales rare, most families obtain land through inheritance. Although the Hindu Succession Act (1956) specifies that land should be divided equally among surviving sons, the law can be circumvented by a will that expresses the head's preferences. Equal division is neither the norm nor the law, and adult sons have incentive to alter their behavior to get larger shares of land. If a head has only daughters, then the land passes from the head's family to the daughter's husband's family and leaves the lineage. Thus, the household head might make land bequest decisions after observing the number of sons that his sons have, since bequeathing land to a son with many daughters and few sons increases the probability that land will eventually leave the lineage. Claimants anticipate the head's preferences and simultaneously make fertility choices to maximize their expected inheritance, taking into account expectations of other claimants' fertility choices. Thus, a claimant has greater incentive to try to have another child when the other claimants have more boys, while remaining indifferent to the other claimants' girls, a prediction I term "strategic fertility". An implication of this fertility pattern is that the average girl in a joint family has more siblings than the average boy, which has been shown to lead to worse health and survival outcomes even when parents' total resources are equitably distributed between their sons and daughters.

In this paper, I test for strategic bequest and fertility behavior as well as the demographic implications of the hypothesis using a nationally representative panel dataset of rural households in India. The results confirm that household heads bequeath a larger share of the land to claimants with more sons. In response, claimants in joint families increase fertility in a race for boys motivated by a desire to increase their inheritance. I find that an additional son among the other claimants in the joint family increases the probability that a claimant will report a pregnancy by 0.8 percent per year. This result is robust to two tests. First, this fertility response is significant in joint families while the head is alive and the land has not yet been distributed, but not in the independent families formed after the head's death. Second, claimants respond to each others' fertility if the joint family head owns land, but not otherwise, suggesting that land bequests are motivating strategic fertility.

As a result of strategic fertility behavior, the average girl who is born in a joint family with two or more claimants has nearly twice as many excess siblings compared to the average girl who is born in a multigenerational family with a single claimant. I calculate that approximately 7 percent of excess female mortality among joint families in Haryana and Rajasthan can be explained by this model. These results suggest a large, but as yet unexamined role for household structure in explaining fertility behavior and poorer outcomes for girls. Thus, this paper contributes to an emerging literature that recognizes the different forms of non-unitary households and family structures observed in developing countries. The joint family literature in particular is sparse, and this paper is one of few papers that incorporates inter and intra-generational dynamics within such families (Rosenzweig and Wolpin 1985; Foster and Rosenzweig 2002; Joshi and Sinha 2003; Edlund and Rahman 2005).

This paper also adds to the strategic bequest behavior literature pioneered by Bernheim, Shleifer, and Summers (1985). Since land bequests form a major share of wealth acquisition in agricultural societies, this framework is particularly useful in understanding behavior in families in rural India. With agricultural land bequests driving differential fertility behavior, Bernheim, Shleifer, and Summers (1985) would suggest that sex differences would increase with the value of land, although this effect might be mitigated by the shift away from farming to other professions.

This paper is organized as follows. Section 2 develops a theoretical model of bequest and fertility behavior in joint families and proposes testable hypotheses. Section 3 describes the data, whereas Section 4 specifies the econometric tests and reports the results. Section 5 concludes with discussion of the results.

2 Theory

In Bernheim, Shleifer, and Summers (1985), parents use bequests to induce children to bring their behavior in line with the parents' preferences. The formulation in this section makes two basic assumptions while adapting that model to the case of farm-based societies in developing countries. First, for reasons outlined earlier, land sales do not occur, so parents do not have the option of selling land and consuming or bequeathing the proceeds. Second, adult daughters leave the household upon marriage to live with their husband's family whereas adult sons may continue to live with or near the parents. This section examines the implications of these two assumptions on the household head's bequest and children's fertility behavior, and shows that fertility behavior may lead to systematic differences in the types of households that girls and boys live in, as a possible explanation for the sex discrimination puzzle. The modeling exercise yields theoretical predictions that can be directly tested in the data.

2.1 Model of fertility choice

This section presents a formal model of bequests with endogenous fertility behavior in joint families. The modeling exercise develops a link between land bequests and fertility, which may in turn influence health and survival outcomes for girls. The theoretical model generates clear predictions that subsequent sections will test empirically.

The family patriarch is the head of the joint family. The head's adult sons are claimants to the family public and private goods while the head is alive, and to the family land once the head is dead. Allocations to each claimant are based on the claimant's family structure. In each period, claimants choose whether to try to have a child or not. Claimants choose the best strategy to maximize their payoff, given the choices made by all other claimants. The head then observes the claimants' family structure and fertility decisions and makes bequest and consumption allocation decisions that maximizes his objective function. I assume that the family head prefers to bequeath land to claimants with more sons in order to perpetuate land ownership within the same lineage. If the head bequeaths land to claimants with only daughters, then that land will leave the family. More land to claimants with more sons implies a greater probability of not having all daughters in the subsequent generation. Assuming no information constraints within the joint family, claimants work recursively to solve the head's problem. Fertility is thus endogenous to bequest and consumption shares.

Consider a single period problem of a family with a head H and claimants indexed by $i \in \{1, \ldots, N\}$. The number of sons and daughters that claimant i has is $\mathbf{n}_i = \{m_i, f_i\}$. The number of boys and girls for all claimants at any point can be written as

$$\mathbf{m}' = [m_1 \ldots m_N] \text{ and } \mathbf{f}' = [f_1 \ldots f_N]$$

Let $\{\mathbf{m}^0, \mathbf{f}^0\}$ represent the number of boys and girls for all claimants at the beginning of a period. $\phi_i \in \{0, 1\}$ represents claimant i's fertility decision in the period, where $\phi_i = 1$ if the claimant reports a pregnancy and 0 otherwise. The fertility decisions made by the set of all claimants is

$$\phi' = [\phi_1 \ldots \phi_N]$$

In this model, the family head determines the bequest share and intrahousehold allocation of private consumption goods for all claimants, as well as the household public good z. The bequest share (κ) and consumption allocation (μ) can be written as follows:

$$\kappa = [\kappa_1 \ldots \kappa_N] \text{ and } \mu = [\mu_1 \ldots \mu_N]$$

$$\text{where } \kappa_i \geq 0 \text{ , } \mu_i \geq 0, z \geq 0 \text{ and } \sum_i \kappa_i = 1 \text{ , } \sum_i \mu_i = 1 \text{ , } z + \sum_i \mu_i x_i = I \text{ for all } i \tag{1}$$

The head's objective is to maximize the utility from bequests, which consists of the probability that land stays within the lineage, as well as a direct utility from bequest. The claimant's objective is to maximize his consumption, given the preferences of the head and the other claimants. To understand the dynamics of these decisions, consider the following sequence of events.

1. Each claimant observes $\{\mathbf{m}^0, \mathbf{f}^0\}$, with preferences well known within the joint family. He decides whether to try to have a child or not (ϕ_i).

2. The head observes $\{\mathbf{m}^0, \mathbf{f}^0\}$ and the fertility decision (ϕ), but not the outcome, for all claimants. He decides the land allocation (κ) as if he were to die in the current period, as well as the consumption allocation (μ) and the amount of public good (z).

3. The head and all claimants observe outcomes $\{\mathbf{m}, \mathbf{f}\}$ from claimants' fertility decisions, as well as whether the head survives. At the end of the period, they realize utility payoffs based on their decisions.

This sequence of events implies that claimants anticipate the head's decisions and respond accordingly. In the two-stage game, I solve the head's problem first, then determine the claimants' reaction functions to the head's decision.

The head's total utility depends on the utility $u_H(.)$ from giving to each claimant. Therefore, the head's problem can be written succinctly as:

$$\max_{\kappa,\mu,z} U_H = \sum_i u_H(\pi_i, \kappa_i, \mu_i, z) \tag{2}$$

where z is the household public good, κ_i is claimant i's bequest share and μ_i is claimant i's consumption allocation. $\pi_i = \pi(m_i)$ is the probability that land bequeathed to claimant i stays within the family lineage such that

$$\frac{\partial \pi_i}{\partial m_i} > 0 \text{ and } \pi_i(m_i^0) > 0 \text{ for all } m_i^0 \tag{3}$$

$\{m_i, f_i\}$ is the outcome of the claimant's fertility decision. This formulation assumes that the head draws direct utility from the act of dividing bequests and consumption allocations among various claimants. He also draws utility from his own consumption of a household public good. The maximization problem is subject to the constraints listed in (1). Solving the problem for all claimants yields the following reaction functions.

$$\kappa_i = \kappa(\mathbf{m}),\ \mu_i = \mu(\mathbf{m}) \text{ and } z = z(\mathbf{m}) \tag{4}$$

The head's preference for bequeathing larger shares of land to claimants with more sons implies

$$\frac{\partial \kappa_i}{\partial m_i} \geq 0 \text{ and } \frac{\partial \kappa_i}{\partial m_{-i}} \leq 0 \tag{5}$$

I term the comparative static in (5) as "strategic bequests", and will directly test for this relationship in the data.[3]

The claimant's expected utility depends on his consumption at the end of the period. Thus, the claimant's objective can be written as

$$\max_{\phi_i} EU_i(x_i, x_i^{\delta}, \mathbf{n}_i) \tag{6}$$

where expectations are taken over the probability that the head survives in the current period. $x_i = x_i(\mathbf{n}, \mu, z)$ is consumption if the head survives and $x_i^{\delta} = x_i^{\delta}(\mathbf{n}, \kappa)$ is the consumption if he dies. In both cases, consumption depends on the number of children the claimant has, since more children are a cost for the claimant. Before the head's death, the claimant's consumption also depends on his share of the household's private (μ) and public resources (z). After the head's death, a claimant's consumption depends on the agricultural output from inherited land (κ). In addition, the claimant draws direct utility from his children ($\mathbf{n}_i$).[4]

[3]One concern might be that the head will decide to grant a larger share of the household consumption goods to claimants who have fewer sons, counter to the result in equation (5). In Appendix A, I show that the qualitative impact of more sons on the claimant's share of consumption goods is the same as the impact on the bequest share.

[4]The claimant can draw utility from current consumption while foreseeing his future role as a household head if

In this specification, fertility choice ϕ_i does not enter directly into the claimant's utility function. To understand how ϕ_i influences $\mathbf{n}_i$, consider that a claimant cannot be sure of the outcome of his fertility decision. He might have a child when he does not want to and might not have a child when he does. The outcome from a fertility decision is

$$m_i = m_i^0 + \mathbf{I}\{\tilde{y} < p\}\phi_i + \tilde{\epsilon}_i^{\phi,m^0} \tag{7}$$

$$f_i = f_i^0 + \mathbf{I}\{\tilde{y} > p\}\phi_i + \tilde{\epsilon}_i^{\phi,f^0} \tag{8}$$

where $\tilde{y}$ is a continuous random variable with distribution $U[0,1]$ and p is the exogenous probability of having a boy. $\tilde{y} < p$ implies that $\mathbf{I}\{\tilde{y} < p\} = 1$ and the claimant has another boy if $\phi_i = 1$. Conversely, $\tilde{y} > p$ implies that $\mathbf{I}\{\tilde{y} > p\} = 1$ and the claimant has a girl if $\phi_i = 1$. $\tilde{\epsilon}_i^{\phi,m^0} \in \{0,1\}$ is a discrete fertility shock whose distribution depends on ϕ_i and m_i^0. Similarly, the distribution of $\tilde{\epsilon}_i^{\phi,f^0} \in \{0,1\}$ depends on ϕ_i and f_i^0. $\tilde{\epsilon}_i = -1$ can represent the loss of a child when no pregnancy is reported, or a still birth when one is. $\tilde{\epsilon}_i = 0$ implies that the claimant has a child if desired. With $\tilde{\epsilon}_i = 1$ and $\phi_i = 1$, twins are born when the claimant reports a pregnancy.

Plugging in the head's reaction functions into all the claimants' problems yields the following solution.

$$\phi_i^* = \phi_i(\mathbf{m}^0) \tag{9}$$

$$\phi_{-i}^* = \phi_{-i}(\mathbf{m}^0) \tag{10}$$

In order to characterize this solution, I impose further restrictions on the preferences claimants' and head's preferences in the next section.

his comprehensive utility consists of two separable parts - utility from consumption as a claimant and utility from bequests as a head.

2.2 Impact of fertility on family structure

Strategic bequests that lead to more pregnancies do not by themselves imply unequal gender-based outcomes. This section illustrates the demographic implications of strategic bequests on the differences in resource allocation between sons and daughters. I link endogenous fertility behavior with poorer outcomes for girls in joint families, even when claimants themselves do not have a preference for boys over girls. To do so, I make standard assumptions on the form of the head's and claimants' utility functions. Assume that the head exhibits declining marginal utility in the bequest share to each claimant and the claimants exhibit declining marginal utility in consumption. These assumptions help to characterize the solution presented in equations (9) and (10).

$$\frac{\partial U_H}{\partial \kappa_i} \geq 0, \frac{\partial^2 U_H}{\partial \kappa_i^2} < 0, \frac{\partial U_i}{\partial x_i} \geq 0, \frac{\partial^2 U_i}{\partial x_i^2} < 0 \tag{11}$$

where x represents the claimant's consumption of household goods as well as children. I further assume that there exist $\hat{m}_i$ and $\hat{f}_i$ such that the marginal utility of an additional child is negative. These conditions are important to rule out situations where a claimant always gains from having an additional child. Thus, given declining benefits from an additional child, a claimant will be observed to have higher probability of trying for another child the fewer sons he already has, or the more sons the other claimants have.

$$Pr\{\phi_i = 1 \mid m_i^0, f_i^0, m_{-i}^0, f_{-i}^0\} > Pr\{\phi_i = 1 \mid m_i^0 + 1, f_i^0, m_{-i}^0, f_{-i}^0\} \tag{12}$$

$$Pr\{\phi_i = 1 \mid m_i^0, f_i^0, m_{-i}^0 + 1, f_{-i}^0\} > Pr\{\phi_i = 1 \mid m_i^0, f_i^0, m_{-i}^0, f_{-i}^0\} \tag{13}$$

I term the theoretical prediction in equation (13) "strategic fertility". The theory also predicts that a claimant will also have lower probability of trying for another child with more own daughters, although the strength of this effect will be less than the impact of own sons (equation 12). In a symmetric problem, this behavior ought to extend to other claimants as well. Therefore, more daughters for other claimants implies that those claimants are less likely to try for another son, reducing a claimant's incentive to try for another child.

$$Pr\{\phi_i = 1 \mid m_i^0, f_i^0, m_{-i}^0, f_{-i}^0\} > Pr\{\phi_i = 1 \mid m_i^0, f_i^0 + 1, m_{-i}^0, f_{-i}^0\} \tag{14}$$

$$Pr\{\phi_i = 1 \mid m_i^0, f_i^0, m_{-i}^0, f_{-i}^0\} > Pr\{\phi_i = 1 \mid m_i^0, f_i^0, m_{-i}^0, f_{-i}^0 + 1\} \tag{15}$$

Suppose two claimants, A and B, with the same initial number of sons and daughters ($m_A^0 = m_B^0, f_A^0 = f_B^0$) have a son and a daughter ($m_A = m_A^0 + 1$ and $f_B = f_B^0 + 1$), respectively. Then the results in (12) and (13) imply that B, who has a new daughter, has greater incentive than A to have another child.

$$Pr\{\phi_B = 1 \mid m_A, m_B\} > Pr\{\phi_A = 1 \mid m_A, m_B\} \tag{16}$$

Without loss of generality, I assume that $m_A^0 = m_B^0 = 0$ and $f_A^0 = f_B^0 = 0$ and that the probability of a pregnancy resulting in a son or daughter is $1/2$. Then

$$\text{No. of siblings for average girl} = \frac{\frac{1}{2}Pr\{\phi_A = 1\} + \frac{3}{2}Pr\{\phi_B = 1\}}{1 + \frac{1}{2}Pr\{\phi_A = 1\} + \frac{1}{2}Pr\{\phi_B = 1\}} \tag{17}$$

$$\text{No. of siblings for average boy} = \frac{\frac{3}{2}Pr\{\phi_A = 1\} + \frac{1}{2}Pr\{\phi_B = 1\}}{1 + \frac{1}{2}Pr\{\phi_A = 1\} + \frac{1}{2}Pr\{\phi_B = 1\}} \tag{18}$$

$$\frac{\text{No. of siblings for average girl}}{\text{No. of siblings for average boy}} = \frac{\frac{1}{2}Pr\{\phi_A = 1\} + \frac{3}{2}Pr\{\phi_B = 1\}}{\frac{3}{2}Pr\{\phi_A = 1\} + \frac{1}{2}Pr\{\phi_B = 1\}} > 1 \tag{19}$$

Similarly, the impact of strategic fertility will imply that the average girl will be observed to have more siblings than the average boy in the aggregate data.

$$E(\text{No. of siblings for average girl}) > E(\text{No. of siblings for average boy}) \tag{20}$$

As a result, the average girl will have systematically more siblings than the average boy to share her resources. The average household resources available to her will be lower even if families are otherwise the same. Therefore, even if a claimant does not discriminate among his children on the basis of gender, the average girl will receive fewer resources than the average boy, and realize poorer health and survival outcomes.

3 Rural Economic and Demographic Survey

Testing the theoretical predictions from Section 2 requires panel or retrospective data that records land inheritance, family structure and fertility decisions as well as other factors that impact inheritance and fertility decisions. The National Council for Applied Economic Research (NCAER) administered the Additional Rural Incomes Survey (ARIS) in 1970-71 to 4,527 households in 259 villages selected from 16 major states of India. Following up on ARIS, NCAER conducted the Rural Economic and Demographic Survey (REDS) among the same households in 1981-82 and 1998-99 (Foster and Rosenzweig 2003). The first wave of REDS in 1981-82 surveyed 250 villages and 4,979 households, excluding nine villages in the state of Assam from the ARIS sample due to a violent insurgency. The second wave of REDS in 1998-99 surveyed 7,474 households consisting of surviving households from the 1981-82 wave, separated households residing in the same village and households from 1970-71 that were missing from the 1981-82 wave. In 1998-99, the REDS sample did not include eight villages that were located in Jammu and Kashmir, where a violent separatist movement perhaps made the survey difficult.[5]

I use data from the 1998-99 wave to test the theory presented in Section 2. My classification differentiates between "families" and "households" since for bequest and inheritance purposes, a split-off household remains within the family, and is not considered an independent family till the head of the previous joint or stem family dies. Tracing the original family of each household

[5]Since the separatist movements in Assam or Jammu and Kashmir are unlikely to be related to family dynamics, the missing villages are not likely to be a source of non-random attrition in the sample. A common source of non-random attrition in panel surveys is from changing household composition due to splitting. The REDS survey tracks split-off family members who were part of the original household in either 1970-71 or 1981-82 and continue to live in the same village, and therefore changing household composition is not a source of bias in the sample.

using the 1981 wave and using information on the circumstances under which the claimant split off, I can categorize households as part of either joint, stem or independent families. I will test the theoretical model using the sample of joint families, while using stem families as a comparison set. Thus, households that were added into the survey for the first time in 1998-99 must be excluded since I cannot determine whether they have been independent since 1981, or are split off members from a joint family household. This leaves 6,203 unique households in 1998-99 originating from 4,026 randomly selected households in the 1981-82 survey.

The survey was administered to three groups of respondents – the household head, every woman in the household between age 15 and 49, and the village head or administrative officer. Household heads answered the economic questionnaire on household migration, formation, division and current structure. They reported why the household split away from the previous household, which is important to determine whether the household is independent or part of a larger joint family. The head also provided detailed information on the source and extent of land holdings, which allowed me to observe how the inheritance was divided by the previous household head. Women in the household between age 15 and 49 answered the demographic questionnaire on pregnancy history, details on each birth, and knowledge and use of contraception. Married women were linked to their husbands who are either family heads or claimants.

I recover an annual retrospective panel dataset from a single wave of observations in 1998-99 since respondents report dates associated with events such as births, deaths and household division. This dataset contains a detailed fertility history for each woman that records whether or not the claimant reported a pregnancy in each year, and the number of living children in that year. Thus, even though the REDS data is not collected annually, it has sufficient historical data for estimating a regression model.

Using the 1998-99 wave of the REDS survey, I construct two datasets. The first is a bequest dataset that contains information on the bequests of land received by 1999 heads from their fathers upon the father's death, and is used to test for strategic bequests in Section 4.1. The second is a "fertility dataset" that contains information on the fertility choices made by the 1999 claimants when the head is still alive, and is used to test for strategic fertility in Section 4.2.

Figure 3 shows four generations of a joint family. The bequest dataset contains the first generation as the head, and the second generation as claimants. In the fertility dataset, the heads are the second generation, and the third generation are claimants. This configuration allows me to test, using the same families, the implications on the previous generation's bequest behavior on the subsequent generation's fertility behavior.

4 Empirical Analysis

The theoretical model of strategic bequests predicts a differential impact of bequest behavior on survival and health outcomes for girls compared to boys. Hence, the first objective of the econometric exercise is to confirm the strategic bequest behavior predicted by equation (5), i.e., whether a claimant's share of the bequest is influenced positively by the number of sons. This establishes the value of sons to claimants in the bequest game.[6] The second objective is to test strategic fertility behavior predicted in equations (13) and (15), i.e., whether a claimant's fertility in a joint family is impacted by the number of boys and girls that the other claimants have. To test this behavior, I propose a "within-family fertility" test that estimates the differential fertility response of claimants in a joint family to the other claimants' boys and girls, a test comparing claimants' behavior when they are in a joint families versus independent families, and a "land ownership" test of claimants' behavior in land owning and non-land owning households. In addition, I calculate the number of siblings born to girls and boys in joint families, and compare this to outcomes in stem families where there is no bequest game. Finally, I calculate the impact of strategic fertility on gender differences in child mortality in six states of India.

[6] Strategic allocations of household public and private consumption goods are not tested since these are not observed in the REDS data.

4.1 Strategic bequests

The bequest dataset contains a cross-sectional snapshot of the family at the time of the head's death. It consists of those land-owning families that were part of a single land-owning household in 1981-82, but had split into at least two households by 1998 following the head's death in the interim.[7] Using the demographic questionnaire, I construct a complete fertility history between waves and calculate the number of sons and daughters for each claimant at the time of the head's death.

Table 1 contains summary statistics from the bequests dataset. The bequest dataset contains 1,266 claimants from 464 heads, with 2.73 claimants per head. Data on the head's characteristics is sparse because all heads had died by the time of the 1998-99 wave and were not directly surveyed. The average size of land inheritance is 1.50 hectares per claimant. Note that each claimant has, on average, 1.1 sons but only 0.9 daughters.

A test for strategic bequest behavior examines how the share of a claimant inheritance (κ_{ij}) varies with the number of sons and daughters ($\mathbf{n_{ij}}$) that claimant i in family j has at the time of the head's death, compared to the sum of the other claimants' sons and daughters ($\sum_{k\neq i}\mathbf{n_{kj}}$). Therefore, I specify the following model.

$$\kappa_{ij} = \alpha_0 + \alpha_1\mathbf{n_{ij}} + \alpha_2\sum_{k\neq i}\mathbf{n_{kj}} + \alpha_3\mathbf{n_{ij}} * r_{ij} + \alpha_4 r_{ij} * \sum_{k\neq i}\mathbf{n_{kj}} + \alpha_5\mathbf{X_{ij}} + \alpha_6\mathbf{Y_j} + \xi_{ij} \tag{21}$$

where $\mathbf{n_{ij}} = [m_{ij}\ f_{ij}]'$, $\alpha_1 = [\alpha_{1m}\ \alpha_{1f}]$, $\alpha_2 = [\alpha_{2m}\ \alpha_{2f}]$, $\alpha_3 = [\alpha_{3m}\ \alpha_{3f}]$ and $\alpha_4 = [\alpha_{4m}\ \alpha_{4f}]$. To confirm the strategic bequest hypothesis, I expect that ($\alpha_{1m} > 0$) and ($\alpha_{2m} < 0$) corresponding to the theoretical predictions in equation (5). The coefficients on two interaction terms $\mathbf{n_{ij}} * r_{ij}$ and $\sum_{k\neq i}\mathbf{n_{kj}} * r_{ij}$ indicate the marginal impact of the number of sons and daughters for a claimant who has moved away from the head's household.[8]

[7]Note that the dataset does not report intended bequest shares while the head is still alive, only the actual shares once he dies. This might create bias if heads' preferences change systematically as they get older. However, if the head's primary objective is to preserve lineage, or if future change in preferences is anticipated by claimants, then I expect this bias to be small.

[8]The impact of moving away is theoretically ambiguous because splitting from the head's household might indicate

This specification must be qualified by controlling for the claimant's residence choice r_{ij} and other observed claimant-specific factors $\mathbf{X_{ij}}$ that might impact bequest preferences. $\mathbf{X_{ij}}$ consists of claimant-specific characteristics such as age at the time of inheritance, years of schooling for the claimant and his wife, which are important since a better educated and technologically savvy claimant might have better access to reproductive technology (including sex-selective abortion). Also included are dummy variables that indicate whether or not the claimant is a farmer and if the claimant's wife works outside the home since these might potentially influence the bequest share. $\mathbf{Y_j}$ includes family specific factors such as the head's education, occupation as farmer and other demographic characteristics. Finally, ξ_{ij} captures unobserved claimant specific factors such as diligence at work or filial relationship with the head, and is assumed to be distributed i.i.d. normal with zero mean.

The results from specification (21), where I test for the influence of family structure on received bequest shares, are presented in Table 4. Column I shows the results of a dual-censored tobit model, where the bequest shares are censored below 0 and above 1. Column II shows the results from an OLS regression. Both sets of estimates are close to each other, and show that an additional son increases the claimant's share of the land bequest by 1 percent. This mirrors the increase in bequest share for the other claimants when they have an additional son (1.2 percent). The opposite effects of relatively equal magnitude indicate that heads bequeath land to claimants with more sons, and that grandsons from different claimants are substitutes for each other. A claimant's birth order has a large influence on the bequest share received by a claimant. An improvement of one position in the birth order increases the bequest share by 8 percent.[9]

Note that a claimant's residence away from the head's household (while remaining in the same family) does not seem to impact his inheritance. While α_3 and α_4 are comparable to α_1 and α_2 in magnitude, the associated standard errors are large and the coefficient cannot be statistically distinguished from zero. Hence, it is unlikely that claimants make fertility and residence choices

that the claimant has been disinherited and is no longer a part of the bequest game, or that the claimant is already in a strong position, irrespective of the number of sons, to receive a significant share of the inheritance.

[9]A closer examination of claimant birth order effects is outside the current model, but could reflect greater certainty about an older claimant's fertility outcomes.

concurrently in order to receive a larger inheritance.[10]

One limitation of this specification is that dependent variables are not independent across observations. Specifically, in a sample where a claimant is the unit of observation, the bequest of one claimant is simply the residual share from the other claimants. Therefore, I separately estimate equation (21) for first-born and other claimants since shares will not be correlated in a sample that includes only first-born claimants, and present these results in Table 5. However, that approach does not use the information that claimant shares must add to one. So I also estimate a multinomial logit model of the head's choices in distribution of land to claimants. I implement this model with only two claimants and seven choices that represent ranges in which the bequest share might fall. Due to small number of observations, I cannot estimate a similar model for families with more than two claimants, or with more categories. The results of the multinomial logit estimation for two-claimant joint families are presented in Table 6. Both the robustness checks are consistent with the strategic bequest results presented earlier.

These results establish that the number of own and other claimants sons are important factors determining the bequest received by the claimant, and provide support to an important assumption made in the theoretical model that sons, not daughters receive the dominant share of land bequests. Thus, claimants have an important incentive to maximize the number of sons they have if they live in a joint family where the head is still alive and owns land.

4.2 Strategic fertility

In the fertility dataset, each observation consists of a man who is older than 15 years of age. Each adult man is counted as one among multiple claimants in a joint family where the head is still alive, as the sole claimant in a stem family where the head is still alive, or else as the head of a nuclear family. In joint and stem families, the claimant can either be co-resident within the same household, or part of a separate household while remaining in the same family.

[10] This suggests that the results from an estimation of fertility choice in equation (4.3) should be similar to the joint bivariate estimation of fertility and residence choice discussed in Appendix B.

The man's wife answers questions on her fertility history, which allows me to create a retrospective panel dataset. Schultz (1972) reports that recalled data on pre and post natal child mortality is more reliable closer to the survey period.[11] Therefore, the sample is restricted to the 1992-98 time period which leaves 43,612 claimant-family-year observations in the panel from 5,090 families over seven years.

In most datasets, the potentially endogenous selection of claimants into classification as joint, stem or independent (nuclear) families is a major concern. The long time period over which the panel is observed, 1981-1999, helps to alleviate some of these concerns. Claimants might live within the household occupied by the joint family head or set up a separate household. Consistent with observed bequest behavior, split off sons retain status as claimants in the joint family household headed by their father as long as the father is alive. Thus, in the dataset used in this paper, co-residence is not a condition for membership in a joint family. Co-residence is a characteristic of a claimant and accounted for using an indicator variable, $r_i \in \{0, 1\}$, that represents the claimant's residence within or outside the head's household respectively. This variable is assigned based on the circumstances of departure and household division as reported in the REDS dataset. Sons who become household heads after their father's death are categorized as independent heads, whereas those who split before their father's death are categorized as part of the joint family till the head dies. Unless a joint family head dies and distributes the bequest, a claimant who lives separately is not classified as an independent (nuclear) family. Thus, the potentially endogenous selection of claimants into joint, stem or independent families is accounted for in case of splits since 1981, which controls for most cases for the analysis in the period 1992-98.

With this assignment, the fertility dataset has 16,162 observations as nuclear families, 7,912 observations in stem families and 19,538 observations in joint families. Table 2 reports the number of claimants in each family type by year. The numbers change over time due to two reasons. First,

[11] Recalled fertility data suffers from bias from two main sources (Schultz 1972). The primary reason is that events in the distant past are reported less frequently than events in the recent past. The secondary reason is that women who are reside in the household in the distant past might be different from those who reside in the household in the recent past. Maternal mortality is a significant factor in the high death rate among adult women in South Asia. Therefore, the mortality rate is higher among more fertile women, leading to non-random sample selection if we survey only women who are alive in 1998-99.

the sample grows as new claimants attain 15 years of age. Second, the number of joint families decreases and the number of independent families increases as heads die and claimants form their own independent families as a result. I assume that both these events occur exogenously.

Table 3 reports the summary statistics for the fertility dataset. Independent couples have on average more children (3.21) than claimants in joint families (2.07). This might reflect the fact that independent heads are older, with average age 43.3 years, compared to claimants in stem (27.6 years) and joint families (31.5 years) and are therefore more likely to have completed their fertility. An important feature of joint families is the significantly worse sex ratio. The ratio of girls to boys is 0.816 in joint families, 0.883 in stem families and 0.969 in independent families. Thus, the data suggests that survival of girls is worse in joint families compared to other family types.[12]

4.3 Within-family fertility

In this section, I test whether the probability that a claimant in a joint family tries to have another child is positively impacted by the number of boys that the other claimants have, corresponding to the theoretical prediction in equation (13).[13] The other claimants' daughters are not future heirs in the family lineage, and have a negative impact on the claimant's own fertility (equation 15). To test these two propositions, I specify a non-linear model with a binary outcome θ_{ijt} that is 1 if a claimant i in joint family j reports a pregnancy in year t, and 0 otherwise.

$$\theta_{ijt} = \beta_0 + \beta_1 \mathbf{n}_{ijt} + \beta_2 \sum_{k \neq i} \mathbf{n}_{kjt} + \beta_3 \mathbf{X}_{ijt} + \beta_4 \mathbf{V}_{ij} + \beta_5 r_{ijt} * \mathbf{n}_{ijt} + \beta_6 r_{ijt} * \sum_{k \neq i} \mathbf{n}_{kjt} + year_t + \mu_j + \epsilon_{ijt} \quad (22)$$

where $\mathbf{n_i} = [m_i f_i]'$, $\beta_1 = [\beta_{1m}\ \beta_{1f}]$, $\beta_2 = [\beta_{2m}\ \beta_{2f}]$, $\beta_5 = [\beta_{5m}\ \beta_{5f}]$ and $\beta_6 = [\beta_{6m}\ \beta_{6f}]$. In this model, the claimant reports a pregnancy based on the number of sons and daughters ($\mathbf{n}_{ijt}$) he

[12]Differences in schooling in Table 3 are consistent with younger couples as claimants in joint families, and relatively older couples as independent heads since formal education has expanded considerably in India over the past few decades (The PROBE Team 1999).

[13]The theoretical mechanism presented Section 2 does not make any predictions about stem and independent families since they have single claimants who cannot exhibit strategic fertility. Therefore, I present results for joint families only.

already has. I expect a negative relationship between the number of children and the probability that the claimant will try for one more, i.e. $\beta_{1m} < 0$ and $\beta_{1f} < 0$. Since sons have value in the bequest game while daughters do not, equation (12) predicts that $\beta_{1m} < \beta_{1f}$. Strategic fertility is identified by the components of β_2. In particular, equations (13) and (15) predict that $\beta_{2m} > 0$, $\beta_{2f} < 0$ and $\beta_{2f} \neq \beta_{2m}$. β_5 indicates the impact of the claimant's own sons and daughters if he is living in a split-off household, whereas β_6 indicates the impact of the other claimants sons and daughters when the claimant is split away. As indicated earlier, the theoretical model makes no clear predictions on the direction or size of β_5 and β_6.

One threat to this specification is from omitted variables that might impact fertility. Therefore, I control for observable time-varying characteristics ($\mathbf{X}_{ijt}$) of the claimant and his partner that impact fertility, such as age, marital status and residence choice as well time-invariant characteristics ($\mathbf{V}_{ij}$) such as years of schooling, and participation in the formal work force. I include year dummy variables to account for time-varying factors that impact fertility across all claimants and families, such as availability of food due to variations in nation-wide monsoon rainfall. ϵ_{ijt} represents unobserved factors that might impact fertility, and is clustered at the family level (μ_j).

I report the results from a probit estimator since probit is well-suited to deal with the serially-correlated error terms in panel data (Wooldridge 2002). Column I of Table 7 presents the marginal effects of probit estimates from the specification in equation (22). As expected, the number of own sons and daughters has a large, negative and statistically significant impact on a claimant's fertility. The probability of the claimant's fertility decreases by 6.5 percent with an additional son, and by 2.2 percent with an additional daughter. In contrast, an additional son for the other claimants increases the probability of a pregnancy by 0.85 percent in a year, a result that is significant at the 1 percent level. The other claimants' daughters have a small impact on the claimant's fertility (-0.5 percent) that is statistically indistinguishable from the null. The important condition of $\beta_{2m} \neq \beta_{2f}$ is confirmed by an F-test with an F-stat of 4.92.

I also examine whether fertility is impacted by moving away from the head's household before his death. The coefficients on the interacted variables in Column I show that moving away has a small impact on own fertility and no particular impact on strategic fertility. The decrease in the probability that a claimant has another child in response to another son changes from 6.5 percent for all claimants to 4.6 percent for those who have formed separate households, perhaps due to greater need for sons for agricultural labor or other household activities. In contrast, the impact of own daughters is the same for split-off claimants as co-resident claimants. Notably, there is almost no marginal influence (0.1 percent) of splitting on the marginal fertility impact of the other claimants' sons.

A possible shortcoming of this specification is that fertility decisions might be influenced by factors that are specific to the joint family, rather than just the claimant. To account for this, I exploit the panel characteristics of the dataset and specify a probit random effects model as well as a logit fixed effects model.[14]

$$\theta_{ijt} = \beta_0 + \beta_1 \mathbf{n}_{ijt} + \beta_2 \sum_{k \neq i} \mathbf{n}_{kjt} + \beta_3 \mathbf{X}_{ijt} + \beta_4 \mathbf{V}_{ij} + \beta_5 r_{ijt} * \mathbf{n}_{ijt} + \beta_6 r_{ijt} * \sum_{k \neq i} \mathbf{n}_{kjt} + family_{jt} + year_t + \epsilon_{ijt} \quad (23)$$

In the probit random effects version of this specification, $family_{jt}$ is a random variable that captures possibly omitted joint family characteristics that may be constant over time but vary between claimants, and others that may be fixed between claimants but vary over time. The parameters of interest are the same as equation (22). In the logit fixed effects version, $family_{jt}$ captures possibly omitted characteristics, such as taste for sons, that are constant over time across members of the joint family.

Column II in Table 7 reports the results of the probit random effects model. The results from this model are not different from those in Column I, though they suggest a possible role for own sons in reducing fertility after the claimant has split away from the head's household. In addition, the other claimants' daughters seem to be statistically different from the null at the 10 percent level. However, since the associated point estimate is different than the point estimate on the impact of the

[14]The probit fixed effects model is not specified (Wooldridge 2002).

other claimants' sons at the 1% level, the validity of the strategic fertility hypothesis is maintained. Column III in the same table reports marginal effects from the logit fixed effects estimation.[15] The other claimants' sons have a positive impact on fertility, although the standard errors are large. However, the other claimants' daughters have a large negative impact on fertility, with the critical statistical test of $\beta_{2m} = \beta_{2f}$ rejected by an F-test.

Table 8 and Figure 4 present additional confirmation of strategic fertility by examining the marginal effects of the estimated impact of the other claimants' sons on a claimant's fertility, as a function of the existing family structure. Reading across each row indicates that the probability of a reported pregnancy always increases for a claimant when the other claimants have more sons. This is consistent with the theoretical prediction that gain in bequest share is greater when the other claimants have more sons than when they have fewer sons. In the vertical column, the fertility response to the other claimants' sons is three times greater when a claimant has no sons, compared to three own sons. This result is also consistent with the theoretical prediction that a large number of sons assures the claimant of a significant bequest share, and there are declining marginal bequest returns from additional sons.[16].

4.4 Joint versus independent families

This test employs the death of the previous head during the period of our study as a natural experiment to observe fertility behavior within the same family.[17] According to Caldwell's framework presented in Section 1, a family with a living head plus multiple adult claimants constitutes a joint family, irrespective of co-residence. Once the head dies, the claimants form independent families. Thus, assuming that the head's death is not associated with fertility behavior, the death of the head and the disbursement of the bequest offers a natural experiment to test for strategic fertility be-

[15] A potential shortcoming of this model is that observations where the dependent variable is constant for all time periods are dropped, which is the case for approximately .

[16] A final concern is the claimant's potentially endogenous choice of co-residence within the same household. This concern is addressed in Appendix B.

[17] Since the family-type assignment of claimants in the 1981 dataset is possibly endogenous, across family comparisons are less credible.

havior within the same family. The theory predicts that strategic fertility ought to be salient only while the head is still alive and the claimants are living in joint families, but not so once the head dies, the bequest is distributed and the claimants form independent families. I test this proposition with a probit specification that interacts the variables in equation (22) with an indicator variable for claimants residing in a joint or independent family.

Table 9 reports the results of this test. The probit marginal effects under 'A' represent the impact of family structure on claimants living in joint families before the head's death, and the coefficients in 'B' represent the impact on the same claimants once they have formed independent households after the head's death. As expected, the claimant's own sons and daughters cause large declines in fertility both before and after the head's death, although the results are significant only in the joint family. The point estimates imply that an additional own son decreases the probability of reporting a pregnancy by 3.3 percent in the joint family, but by 13.5 percent in the independent family. The most probable reason for this large difference is that even controlling for age, higher order births are more likely after head's death, and thus the marginal reduction in the probability of an additional pregnancy is greater. The most notable result in Table 9 is that the other claimant's sons have a positive and statistically significant impact on own fertility in the joint family (+ 1.3 percent and statistically significant at the 5 percent level), but has virtually no impact on the same claimants following the head's death once they have formed independent families (+0.6 percent but statistically indistinguishable from the null, with the relevant F-test rejecting $\beta_{2m} = \beta_{2f}$ in joint families but not so in independent families). As predicted by equation (15), the other claimants' daughters have a statistically significant, negative impact on fertility in joint families (-1.3 percent), but a statistically insignificant in independent ones. These results imply that claimants' consider each others' fertility only insofar that the head is alive and has not yet distributed his land, but not so once the head dies, land division is complete and claimants are heads of independent families.

4.5 Land ownership

The final test uses Bernheim, Shleifer and Summer's (1985) prediction that the strategic bequest game is impacted by the size of the bequest. Analogously, I test whether strategic fertility is influenced by the presence of a bequest. If the head of a joint family owns no land that he can bequeath, then claimants have no incentive for strategic fertility. In this case, neither the other claimants' sons nor daughter will be significant in the claimants' fertility decision. However, land ownership should induce a positive fertility response from claimants towards To test this hypothesis, I estimate a probit model where the variables in equation (22) are interacted with indicator variables for landless and land owning family.

Table 10 presents the results of the test of strategic fertility specified in section 4.5. Column A represents claimants in joint families where the head does not own any land. Column B reports claimants in joint families with a land-owning head.

The result is that while the claimant's own sons are the only statistically significant determinants of fertility in joint families that do not own land, both own family structure as well as other claimants' boys are significant in land-owning families. The point estimates imply that an additional son for other claimants increases the fertility rate by 0.82 percent in land owning families. This estimate is close to the 0.85 percent increase in the fertility rate reported in section 4.3. The same coefficient for landless families is larger, i.e., 2.3 percent, but cannot be statistically distinguished from zero. This result suggests that there is greater variation in the response of the claimant's own fertility to other claimants' sons when the head does not own land than when he does. Thus, the results of this test are consistent with the prediction that strategic fertility is primarily a phenomenon among land-owning families, offering an explanation why the sex ratio is relatively worse in such families.

4.6 Implications of strategic fertility

To show how strategic fertility behavior yields health and mortality differences in outcomes for boys and girls, I propose two mechanisms. First, the differences in fertility responses imply that the average girl in the population lives in a family that has systematically more children than the average boy. Thus, even without differences in resource allocations by parents towards children of different gender, the average girl will receive smaller share of resources than the average boy, explaining poorer outcomes. Second, mortality rates are greater, and different for boys and girls, for higher order births. Thus, the joint effect of these two mechanisms implies greater mortality for girls, which I calculate using existing estimates from Arnold, Choe, and Roy (1998).

4.6.1 Excess siblings

To test the first mechanism, I check whether the average girl has more siblings than the average boy. Let f_{ij} and m_{ij} represent the number of daughters and sons born to claimant i in family j. Correspondingly, s_{ij} is the number of siblings for any one of that claimant's children. $\bar{s}_f$ and $\bar{s}_m$ represent the number of siblings for the average girl and boy respectively.

$$\bar{s}_f = \frac{\sum_{i,j}(s_{ij} * f_{ij})}{\sum_{i,j} f_{ij}} \text{ and } \bar{s}_m = \frac{\sum_{i,j}(s_{ij} * m_{ij})}{\sum_{i,j} m_{ij}} \tag{24}$$

The excess number of siblings for the average girl is $\bar{s}_f - \bar{s}_m$. I expect this to be positive, and larger for joint families with multiple claimants than for stem families that have similar observed characteristics (see Table 3), but only a single claimant and hence no strategic fertility.

Table 11 calculates the sibling statistics for stem and joint families. $\bar{s}_f$ is 2.761 whereas $\bar{s}_m$ is 2.481 in joint families. Hence, the average girl has 0.280 excess siblings compared to the average boy in joint families. Contrast this with 0.156 excess siblings for the average girl in stem families. The difference in the excess siblings between stem and joint families is driven by fewer number of siblings for the average boy in a joint family. $\bar{s}_f$ for stem families (2.757) is close to $\bar{s}_f$ for joint families (2.761). However, the difference in $\bar{s}_m$ for stem families (2.600) and $\bar{s}_m$ for joint families (2.481) is large.[18] This is consistent with the theory presented in Section 2 that predicts that a joint family with many boys is more likely to observe declines in fertility compared to similar stem families, or families with many girls in either family type.

Thus, the results in this section confirm that girls born in joint families live in households that are systematically larger than where boys are born. The comparison with stem families suggests that this is driven by the specific strategic fertility behavior observed in joint families.

4.6.2 Differential mortality

To test the second mechanism, I perform back-of-the-envelope calculations to examine the impact of strategic fertility behavior on differential mortality for girls and boys in six major states of India. I estimate the marginal effect of the claimant's existing sons and daughters and the other claimants' sons ($\mathbf{B}$) and count the number of observations in each cell ($\mathbf{\Delta}$).[19] I assume that the probability of a male birth is $p = \frac{1}{2}$. Arnold, Choe, and Roy (1998) report the mortality rates for boys ($\mathbf{\Omega_m}$) and girls ($\mathbf{\Omega_f}$) given the claimant's existing family structure. Thus, the excess female deaths from strategic fertility behavior is

$$\mathbf{B} \times \mathbf{\Delta} \times (1 - p)\mathbf{\Omega_f} - \mathbf{B} \times \mathbf{\Delta} \times p\mathbf{\Omega_m} \tag{25}$$

I also estimate the probability of another pregnancy reported conditional only on the claimant's sons and daughters ($\mathbf{\Gamma}$). Therefore, the overall excess female deaths from having a child is

$$\mathbf{\Gamma} \times \mathbf{\Delta} \times (1 - p)\mathbf{\Omega_f} - \mathbf{\Gamma} \times \mathbf{\Delta} \times p\mathbf{\Omega_m} \tag{26}$$

Thus, the fraction of excess female deaths in India due to strategic fertility behavior is

[18] Excess siblings for the average girl is not a trivial outcome of a sex ratio skewed against girls. Even with a skewed sex ratio, random assignment of girls and boys to households will yield the same number of siblings for both the average girl and boy.

[19] Following Arnold, Choe, and Roy (1998), the categories corresponding to (m_i, f_i, $\sum_{k\neq i} \mathbf{m}_k$) are $(0,0,0)$, $(0,0,1)$, $(0,0,2)$, $(0,1,0)$, $(0,1,1)$, $(0,1,2)$, $(0,2,0)$, $(0,2,1)$, $(0,2,2)$, $(1,0,0)$, $(1,0,1)$, $(1,0,2)$, $(1,1,0)$, $(1,1,1)$, $(1,1,2)$, $(2,0,0)$, $(2,0,1)$, and $(2,0,2)$.

$$\frac{\mathbf{B} \times \mathbf{\Delta} \times (1-p)\mathbf{\Omega_f} - \mathbf{B} \times \mathbf{\Delta} \times p\mathbf{\Omega_m}}{\mathbf{\Gamma} \times \mathbf{\Delta} \times (1-p)\mathbf{\Omega_f} - \mathbf{\Gamma} \times \mathbf{\Delta} \times p\mathbf{\Omega_m}} \tag{27}$$

Table 12 reports the fraction of excess female deaths explained by strategic fertility behavior for six states.[20] The greatest impact of strategic fertility is in the Green Revolution states of Punjab, Haryana and Rajasthan where the value of land (Figure 1) is greatest, and agriculture yields the largest share of income (Figure 2). Four percent of the female shortage in Punjab, and 7 percent of the shortage in Haryana and Rajasthan is explained by my model. In Orissa, where agricultural yields are low, strategic fertility tied to bequests of agricultural land is not a large incentive for fertility behavior. In Kerala, where matrilineal descent is practiced among many communities, 39 percent of the shortage of *boys* is explained by fertility induced by land bequest motives. These back of the envelope calculations also predict a small shortage of boys in Tamil Nadu, which does not correspond with Census data. One possible explanation for this mismatch is that the fertility rates are calculated for the period 1992-98, whereas the source for the sex ratio is the 2001 Census.

5 Discussion

This paper demonstrated a mechanism by which bequest behavior in land-owning joint families in rural India impacts gender differences in health and survival outcomes. The theoretical model showed that in a patrilocal society, heads will prefer to bequeath land to claimants with more sons in order to preserve land within the family in future generations. This motivates a race for boys among claimants, manifested by strategic fertility, leading to family structures where the average girl has more siblings than the average boy. Even without intra-household differences in allocation, this result implies fewer resources for the average girl. Thus, fairly benign behavior that manifests itself in differential stopping rules has the potential to explain large and near universal differences in health and mortality outcomes.

I test both the strategic bequest and strategic fertility hypotheses. I confirm that heads prefer

[20] Arnold, Choe, and Roy (1998) report mortality rates only for these states.

claimants with more sons, and as a result claimants' fertility behavior responds strategically to the family structures of the other claimants. As expected, this result is more pronounced in land owning families relative to landless families, offering a possible explanation why sex differences are larger in relatively prosperous families. Strategic fertility is salient among claimants in joint families before the head's death and distribution of the bequest compared to the same claimants in independent families once the head has died and the inheritance has been received.

The results should be read with a few caveats. First, strategic fertility does not rule out overtly discriminatory behavior by claimants against girls. Bequests might motivate significant foeticide, infanticide or differences in resource allocation that I do not estimate in the empirical analysis. For example, Jayachandran and Kuziemko (2010) report that mothers shorten the time between pregnancies after a daughter's birth compared to a son's, resulting in a lower breastfeeding and poorer lifelong health outcomes, a result that is entirely consistent with the model presented in this paper. In another paper, Rosenblum (2010) presents a hypothesis based on differential vaccination that is also consistent with my results. Additionally, sex bias might be motivated for reasons other than bequests, such as the asymmetric labor market and cultural returns mentioned earlier. The impact of strategic bequests and fertility are congruent to these reasons, not in opposition to them. Finally, the model relies explicitly on the value of land as a permanent agricultural asset as well as the social institution of women leaving their parents' family at the time of marriage. Therefore, I do not address gender differences in societies where land is not central to the production process, or that have alternative types of social institutions.

From a policy perspective, the results underscore the influence of differential bequest behavior on even apparently benign fertility behavior. Legal changes in the 1980s and 90s in a number of southern Indian states granted daughters inheritance rights to agricultural land if the head dies without a will. Deininger, Goyal, and Nagarajan (2010) report that the bargaining power of daughters in the bequest game improved as a result of these changes, leading to both greater asset ownership as well as improvements in human capital among women in these states. A recent amendment to the Hindu Succession Act (2005) extended these rights nationally, although it is unclear whether the law will motivate greater equality in bequests, or increased writing of wills. Finally, land ownership is a key driver of strategic fertility behavior, which suggests that the shift towards other forms of bequests such as investments in professional education might alleviate an important cause of differential gender outcomes.

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Appendices

A Intra-household allocation

Suppose $\eta(\mathbf{m})$ represents the distribution of power across different claimants in the joint family such that $\sum_i \eta_i(\mathbf{m}) = 1$. If the birth of a son increases a claimant's power within the joint family, then $\frac{\partial \eta(\mathbf{m})}{\partial m_i} > 0$. Thus, in a collective model of efficient intra-household allocation of private and public goods, the household head faces the following optimization problem.

$$\max_{\mu_1,\ldots,\mu_N,z} \sum_i \eta_i(\mathbf{m}) u_i(\mu_i X, z) \tag{28}$$

$$\text{such that } z + \sum_i \mu_i x_i = I \text{ and } \sum_i \mu_i = 1 \tag{29}$$

Then the first order conditions yield

$$\eta_i(\mathbf{m}) \frac{\partial u_i}{\partial \mu_i} = \eta_{-i}(\mathbf{m}) \frac{\partial u_{-i}}{\partial \mu_{-i}} \tag{30}$$

$$\text{or } \frac{\eta_i(\mathbf{m})}{\eta_{-i}(\mathbf{m})} = \frac{MU_{-i}}{MU_i} \tag{31}$$

Assuming that claimants exhibit declining marginal utility in consumption of private goods, this condition implies that an increase in $\eta_i(\mathbf{m})$ due to the birth of a son will result in an increase in the allocation μ_i for the claimant, or

$$\frac{\partial \mu_i}{\partial m_i} \geq 0, \frac{\partial \mu_i}{\partial m_{-i}} \leq 0 \tag{32}$$

B Endogenous residence choice

One concern with the tests of strategic fertility presented in Section 4.2 is the possibly endogenous determination of residence choice with fertility. Claimants are more likely to leave the joint family's household either when public good provision or the possible share in bequest share declines. The results in Section 4.1 suggest that a claimant's residence does not significantly impact his share of the bequest. I check this result by estimating a bivariate normal probit model for the within-family test. The parameters of interest and associated theoretical predictions are the same as in equation (22) respectively.

Column I in Table 13 reports the results from joint determination of fertility and residence choice in the within family test presented in Section 4.3. Column II reports the results from joint determination of fertility and residence choice in the land ownership test (Section 4.5), with coefficients under 'A' representing claimants in joint families where the head does not own any land and coefficients under 'B' representing claimants in joint families with a land-owning head.

The results from this model are not materially different from those in reported in Table 7, confirming that residence choice is not a significant factor in the strategic fertility game.

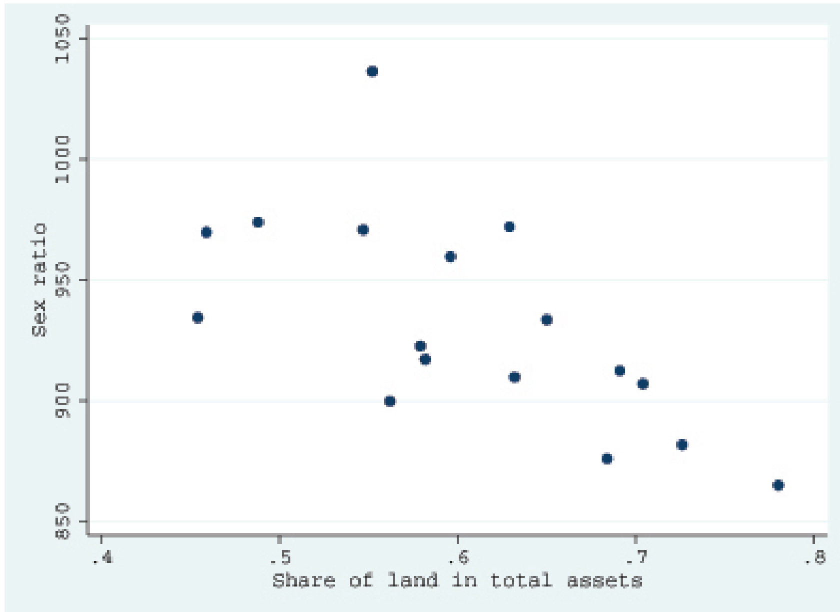


Figure 1: **Importance of land vs. Sex ratio**
Source: Govt. of India (1998) and Census of India (2001).

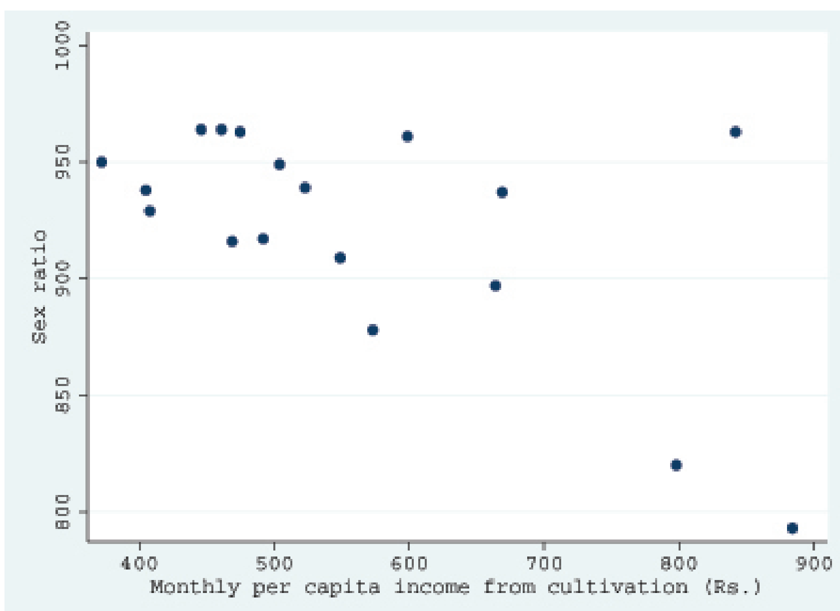


Figure 2: **Agricultural income vs. Sex ratio**
Source: Govt. of India (1998) and Census of India (2001).

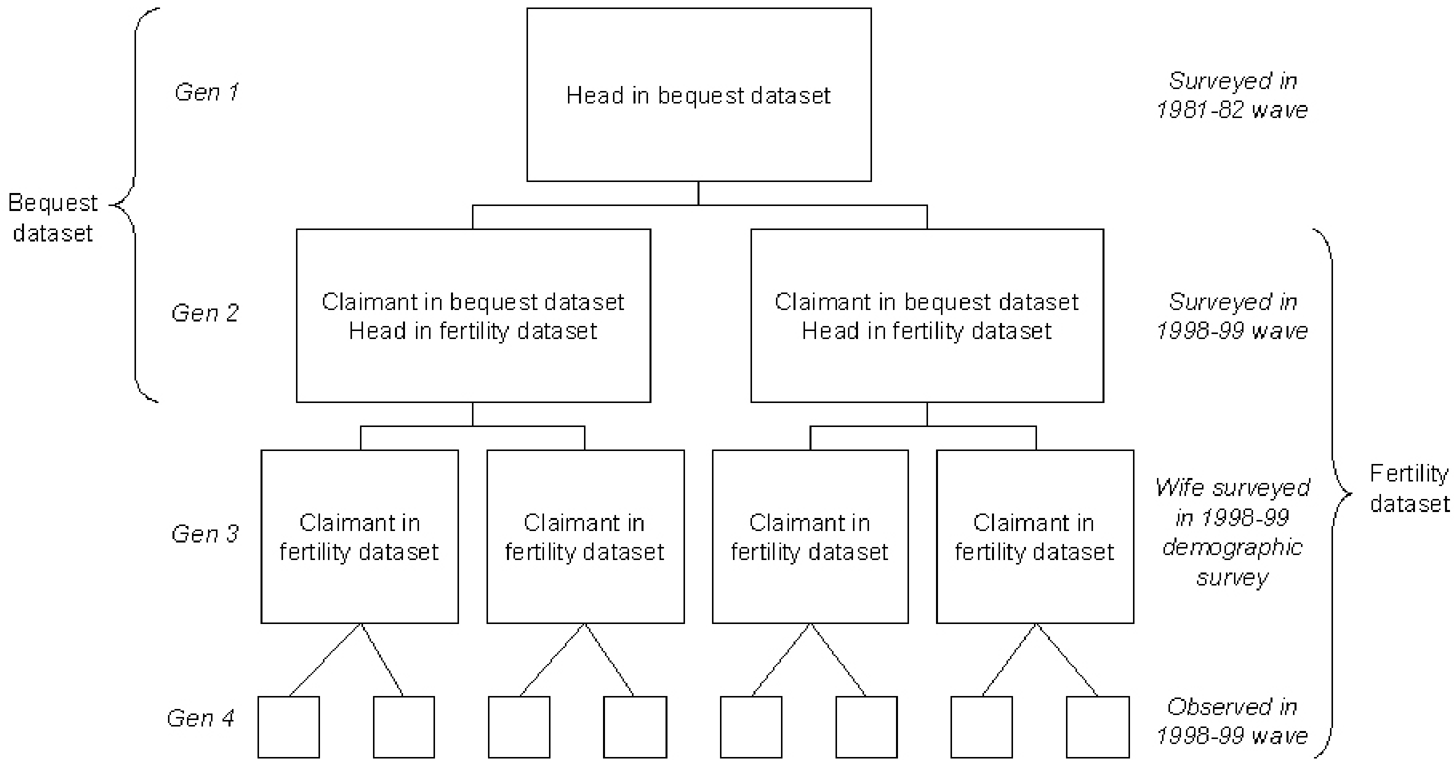


Figure 3: **Four generations of a joint family**

Table 1: **Summary statistics: Bequests dataset**

	Mean (or percent)	Std. dev.
Number of heads	464	
Head's schooling	1.2 years	2.561
Hindu	90.5%	0.29
Brahmin	9.2%	0.29
Other Upper Caste	30.4%	0.46
Scheduled Caste	8.1%	0.27
Claimants per head	2.73	0.97
Claimant's characteristics		
Number of claimants	1,266	
Age	33.9 years	9.8
Size of land inherited	1.50 hectares	1.66
Number of sons	1.1	1.2
Number of daughters	0.9	1.2
Split from head's household	24.2%	0.4
Married	94.6%	0.23
Claimant's schooling	5.8 years	4.9
Wife's schooling	2.9 years	3.9
Occupation as farmer	72.3%	0.45
Wife works outside home	30.6%	0.73

Notes: Variables as reported at time of inheritance. Head's characteristics not available since survey data collected after head's death. Source: REDS 1998-99.

Table 2: **Claimants in fertility dataset**

Year	Independent	Stem	Joint	Total
1992	2,120	1,092	2,934	6,146
1993	2,169	1,115	2,928	6,212
1994	2,227	1,126	2,877	6,230
1995	2,310	1,142	2,800	6,252
1996	2,385	1,145	2,726	6,256
1997	2,453	1,150	2,655	6,258
1998	2,498	1,142	2,618	6,258
Total	**16,162**	**7,912**	**19,538**	**43,612**
Share of Total	37.1%	18.1%	44.8%	100%

Source: REDS 1998-99.

Table 3: **Summary statistics: Fertility dataset**

	Independent	**Stem**	**Joint**	**Total**
N (claimant-family-year)	16,162	7,912	19,538	43,612
Age	43.3 years	27.6 years	31.5 years	35.4 years
	(12.4)	(7.6)	(8.9)	(12.0)
Boys per claimant	1.63	1.00	1.14	1.30
	(1.21)	(1.18)	(1.15)	(1.21)
Girls per claimant	1.58	0.88	0.93	1.17
	(1.39)	(1.14)	(1.13)	(1.28)
Total children	3.21	1.88	2.07	2.47
	(1.87)	(1.87)	(1.77)	(1.92)
Sex ratio (girls/boys)	0.97	0.88	0.82	0.90
Other claimants' boys			1.98	
			(2.30)	
Other claimants' girls			1.61	
			(2.04)	
Split from head's household			28%	
			(0.45)	
Married	86%	78%	83%	83%
	(0.34)	(0.42)	(0.38)	(0.37)
Claimant's schooling	5.5 years	7.1 years	6.8 years	6.4 years
	(4.48)	(4.93)	(4.91)	(4.95)
Age at headship	32.0 years			
	(10.15)			
Woman working outside	35%	29%	27%	30%
	(0.64)	(0.63)	(0.53)	(0.59)
Woman's schooling	3.1 years	4.7 years	3.8 years	3.7 years
	(4.15)	(4.73)	(4.41)	(4.42)
Hindu	90.3%	87.9%	89.0%	89.3%
	(0.30)	(0.33)	(0.31)	(0.31)
Brahmin	7.1%	6.8%	9.4%	8.1%
	(0.26)	(0.25)	(0.29)	(0.27)
Other Upper Caste	25.8%	28.6%	29.1%	27.8%
	(0.44)	(0.45)	(0.45)	(.45)
Scheduled Caste	13.3%	11.8%	10.9%	11.9%
	(0.34)	(0.32)	(0.31)	(0.32)

Note: Value in parentheses is standard deviation. Source: REDS 1998-99.

Table 4: **Strategic bequest results**

Dependent variable: Share of land		
	I: Tobit	**II: OLS**
Number of own sons	**0.010** **	**0.010** **
	(0.005)	(0.005)
Number of own daughters	0.001	0.001
	(0.005)	(0.005)
Other claimants' sons	**-0.012** ***	**-0.012** ***
	(0.003)	(0.003)
Other claimants' daughters	-0.004	-0.004
	(0.003)	(0.003)
Split	-0.020	-0.016
	(0.019)	(0.018)
Number of own sons * Split	0.003	0.001
	(0.009)	(0.009)
Number of own daughters * Split	0.014	0.012
	(0.009)	(0.009)
Other claimants' sons * Split	-0.006	-0.006
	(0.006)	(0.005)
Other claimants' daughters * Split	0.001	0.002
	(0.006)	(0.006)
Birth Order	**-0.080** ***	**-0.079** ***
	(0.006)	(0.006)

Notes: N = 1067 claimants from 430 groups. The number of left censored, uncensored and right censored observations in the dual-censored tobit model is 35, 1019 and 13 respectively. Standard errors are clustered at joint family level. *** $p < 1\%$. ** $p < 5\%$. * $p < 10\%$. Source: REDS 1998-99.

Table 5: **Strategic bequest results by birth order**

Dependent variable: Share of land		
	Birth Order = 1	**Birth Order** > 1
Number of own sons	**0.023** ***	**0.018** ***
	(0.008)	(0.007)
Number of own daughters	0.006	0.010
	(0.008)	(0.007)
Other claimants' sons	**-0.021** ***	**-0.018** ***
	(0.007)	(0.004)
Other claimants' daughters	**-0.017** ***	**-0.009** **
	(0.007)	(0.004)
Split	0.000	-0.016
	(0.028)	(0.028)
Number of own sons * Split	-0.003	0.001
	(0.015)	(0.013)
Number of own daughters * Split	0.008	0.026 *
	(0.014)	(0.015)
Other claimants' sons * Split	-0.008	-0.005
	(0.013)	(0.007)
Other claimants' daughters * Split	0.009	-0.001
	(0.014)	(0.007)

Notes: N = 1067 claimants from 430 groups. Dual-censored random effects tobit model. Standard errors are clustered at joint family level. *** $p < 1\%$. ** $p < 5\%$. * $p < 10\%$. Source: REDS 1998-99.

Table 6: **Multinomial logit results of test of strategic bequest**

Category (Share of Land)	**0 to 0.25**	**0.25 to 0.40**	**0.40 to 0.50**	**0.50 to 0.60**	**0.60 to 0.75**	**0.75 to 1**
Number of own sons	0.601 *	1.020 *	-0.879 *	**1.285** **	0.073	-0.485
	(0.32)	(0.59)	(0.47)	(0.54)	(0.52)	(0.38)
Number of own daughters	-0.038	-0.053	0.131	-0.858	-2.276	-0.049
	(0.35)	(0.47)	(0.33)	(0.62)	(1.40)	(0.38)
Other claimant's sons	-0.453	0.547	0.356	**-2.886** ***	0.769 *	0.129
	(0.34)	(0.53)	(0.28)	(1.03)	(0.46)	(0.30)
Other claimant's daughters	0.012	-0.614	0.399	**1.380** **	-2.106 *	-0.058
	(0.33)	(0.74)	(0.34)	(0.58)	(1.24)	(0.38)
Claimant's age	-0.006	0.065	0.071	-0.054	-0.108	-0.055
	(0.05)	(0.10)	(0.05)	(0.07)	(0.08)	(0.06)
Other claimant's age	0.035	-0.089	-0.097 *	0.151	0.056	0.025
	(0.06)	(0.11)	(0.06)	(0.10)	(0.07)	(0.06)
Birth order	-0.132	-1.591	0.591	0.445	-0.461	-1.096
	(1.02)	(2.64)	(1.13)	(1.69)	(1.43)	(1.10)
Observations in category	12	6	15	10	11	14

Notes: Share of Land = 0.50 is the base outcome. *** $p < 1\%$. ** $p < 5\%$. * $p < 10\%$. Source: REDS 1998-99.

Table 7: **Results of test of strategic fertility within joint families**

Dependent variable: Reported pregnancy			
	I: Probit	**II: Probit RE**	**III: Logit FE**
Number of own sons	**- 0.065** ***	**- 0.067** ***	**-0.025** ***
	(0.007)	(0.026)	(0.006)
Number of own daughters	**- 0.022** ***	**-0.026** ***	**-0.017** ***
	(0.005)	(0.025)	(0.004)
Other claimants' sons	**0.008** ***	**0.008** ***	0.001
	(0.003)	(0.001)	(0.001)
Other claimants' daughters	-0.005	- 0.006 *	**-0.007** ***
	(0.003)	(0.016)	(0.002)
Number of own sons * Split	- 0.046 *	**-0.048** ***	**-0.022** **
	(0.026)	(0.091)	(0.009)
Number of own daughters * Split	- 0.020	-0.019	-0.004
	(0.017)	(0.071)	(0.005)
Other claimants' sons * Split	- 0.001	0.001	-0.005
	(0.009)	(0.043)	(0.004)
Other claimants' daughters * Split	0.015	0.014	0.005
	(0.010)	(0.048)	(0.004)
Number of observations	7,522	7,522	6,287
$\beta_{2m} - \beta_{2f}$	**0.013** **	**0.014** ***	**0.008** ***
F statistic	4.92	6.64	15.07

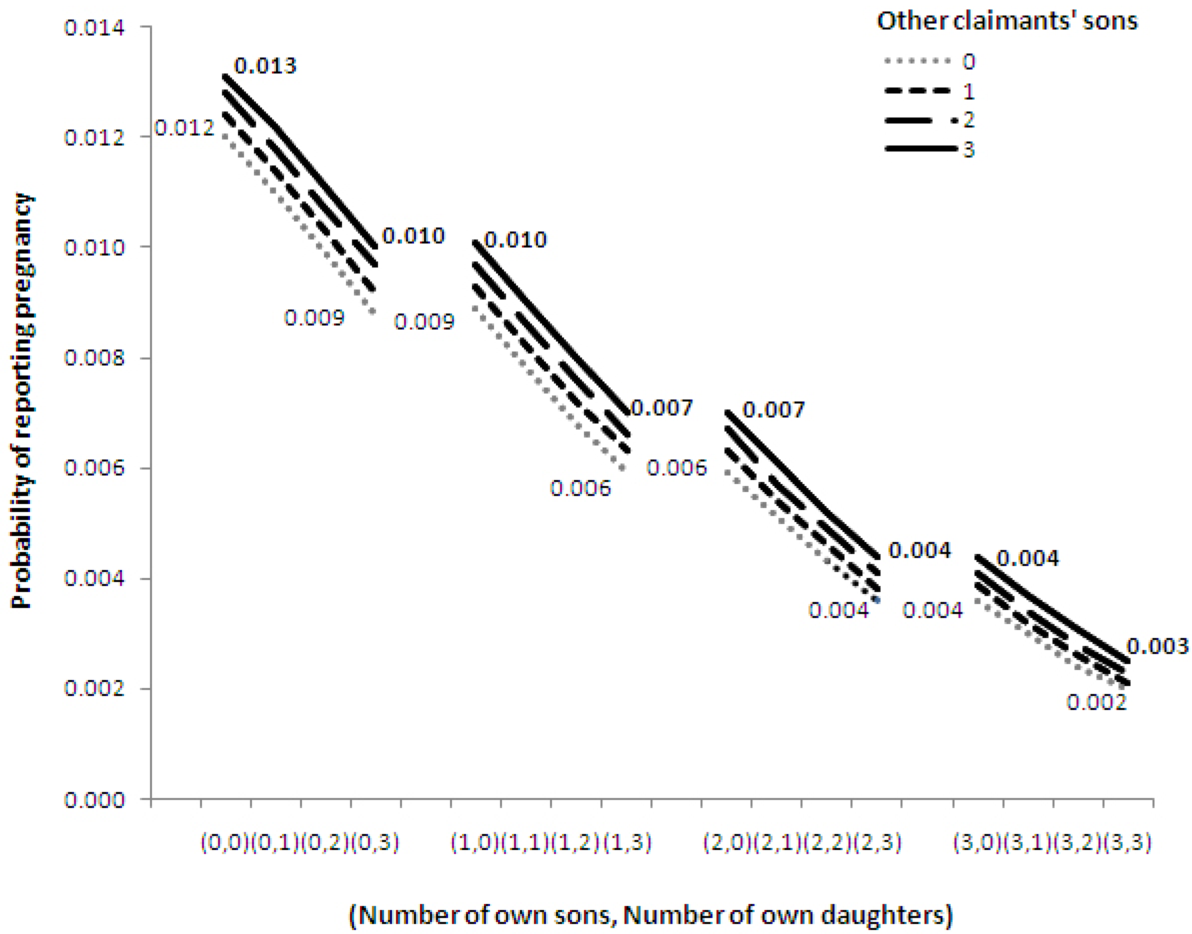
Notes: Values in parentheses are standard errors. Standard errors are clustered at joint family level. *** $p < 1\%$. ** $p < 5\%$. * $p < 10\%$. N = 7,522 in 599 joint families. Source: REDS 1998-99.

Table 8: **Marginal effects of within family test**

Dependent variable: Reported pregnancy					
Number of own sons	**Number of own daughters**	**Other claimants' sons**			
		0	**1**	**2**	**3**
0	0	**0.0120** ***	**0.0124** ***	**0.0128** ***	**0.0131** ***
		(0.0038)	(0.0040)	(0.0043)	(0.0045)
0	1	**0.0110** ***	**0.0114** ***	**0.0118** ***	**0.0122** ***
		(0.0034)	(0.0037)	(0.0039)	(0.0042)
0	2	**0.0099** ***	**0.0103** ***	**0.0107** ***	**0.0111** ***
		(0.0031)	(0.0034)	(0.0036)	(0.0039)
0	3	**0.0088** ***	**0.0092** ***	**0.0097** ***	**0.0100** ***
		(0.0028)	(0.0030)	(0.0033)	(0.0036)
1	0	**0.0089** ***	**0.0093** ***	**0.0097** ***	**0.0101** ***
		(0.0027)	(0.0030)	(0.0032)	(0.0035)
1	1	**0.0078** ***	**0.0082** ***	**0.0086** ***	**0.0090** ***
		(0.0024)	(0.0026)	(0.0029)	(0.0031)
1	2	**0.0068** ***	**0.0072** ***	**0.0076** ***	**0.0080** ***
		(0.0021)	(0.0023)	(0.0026)	(0.0028)
1	3	**0.0059** ***	**0.0063** ***	**0.0066** ***	**0.0070** ***
		(0.0018)	(0.0020)	(0.0023)	(0.0025)
2	0	**0.0059** ***	**0.0063** ***	**0.0067** ***	**0.0070** ***
		(0.0018)	(0.0020)	(0.0022)	(0.0025)
2	1	**0.0051** ***	**0.0054** ***	**0.0057** ***	**0.0061** ***
		(0.0015)	(0.0017)	(0.0019)	(0.0022)
2	2	**0.0043** ***	**0.0046** ***	**0.0049** ***	**0.0052** ***
		(0.0013)	(0.0015)	(0.0017)	(0.0019)
2	3	**0.0036** ***	**0.0038** ***	**0.0041** ***	**0.0044** ***
		(0.0011)	(0.0013)	(0.0014)	(0.0016)
3	0	**0.0036** ***	**0.0039** ***	**0.0041** ***	**0.0044** ***
		(0.0011)	(0.0013)	(0.0015)	(0.0016)
3	1	**0.0030** ***	**0.0032** ***	**0.0034** ***	**0.0037** ***
		(0.0009)	(0.0011)	(0.0012)	(0.0014)
3	2	**0.0024** ***	**0.0026** ***	**0.0028** ***	**0.0031** ***
		(0.0008)	(0.0009)	(0.0010)	(0.0012)
3	3	**0.0020** ***	**0.0021** ***	**0.0023** ***	**0.0025** ***
		(0.0006)	(0.0007)	(0.0009)	(0.0010)

Notes: Standard errors in parentheses are clustered at joint family level. *** $p < 1\%$. Coefficients for marginal effects evaluated at mean of independent variables. Marginal effects for Own sons > 3, Own daughters > 3 or Other claimants' sons > 3 not shown. N = 7,522 in 599 joint families. Source: REDS 1998-99.

Figure 4: **Marginal effects of within family test**



Notes: Coefficients for marginal effects evaluated at mean of independent variables. Marginal effects for Own sons > 3, Own daughters > 3 or Other claimants' sons > 3 not shown. N = 7,522 in 599 joint families. Source: REDS 1998-99.

Table 9: **Strategic fertility in joint versus independent families**

Dependent variable: Reported pregnancy		
	A: Claimant in joint family (Before head's death)	**B: Independent family (After head's death)**
Number of own sons	**- 0.033** **	- 0.135
	(0.017)	(0.272)
Number of own daughters	**- 0.022** **	- 0.519
	(0.012)	(0.481)
Other claimants' sons	**0.013** **	0.006
	(0.005)	(0.031)
Other claimants' daughters	**- 0.013** **	- 0.078
	(0.007)	(0.063)
$\beta_{2m} - \beta_{2f}$	**0.026 ****	0.084
F statistic	5.84	1.32

Notes: Values in parentheses are standard errors. Standard errors are clustered at joint family level. ** $p < 5\%$. N = 768 in 125 joint families. Source: REDS 1998-99.

Table 10: **Strategic fertility in land owning and landless families**

Dependent variable: Reported pregnancy		
	A: Landless head (Marginal Effects)	**B: Landowning head (Marginal Effects)**
Number of sons	**-0.092** ***	**-0.064** ***
	(0.035)	(0.007)
Number of daughters	-0.021	**-0.022** ***
	(0.028)	(0.005)
Other claimants' sons	0.023	**0.008** ***
	(0.021)	(0.003)
Other claimants' daughters	-0.003	-0.004
	(0.022)	(0.004)
$\beta_{2m} - \beta_{2f}$	**0.026 *** **	**0.012 ***
F statistic	7.02	3.32

Notes: Values in parentheses are standard errors. Standard errors clustered at joint family level. *** $p < 1\%$. ** $p < 5\%$. * $p < 10\%$. N = 7,522 in 599 joint families. Source: REDS 1998-99.

Table 11: **Excess siblings for average girl**

Household type	**Male births** $\sum m_{ij}$	**Female births** $\sum f_{ij}$	**Siblings per girl** $\bar{s}_f$	**Siblings per boy** $\bar{s}_m$	**Excess siblings per girl** $\bar{s}_f - \bar{s}_m$
Stem	1,259	1,113	2.757	2.600	0.156
Joint	3,179	2,666	2.761	2.481	0.280

Notes: Independent families have a couple plus minor children. Stem families have family head and one adult claimant. Joint families have family head and two or more adults claimants. Analysis based on 1998 family structure. Source: REDS 1998-99.

Table 12: **Differential mortality due to strategic fertility**

State	Number of observations	Excess female deaths Strategic fertility	All Reasons	Fraction of excess female deaths explained	Observed sex ratio (Census)
Haryana	304	0.021	0.286	7%	861
Rajasthan	471	0.025	0.376	7%	922
Punjab	491	0.010	0.285	4%	874
Orissa	232	0.002	0.161	1%	972
Tamil Nadu	127	-0.003	-0.026	-11%	986
Kerala	186	-0.003	-0.008	-39%	1,058

Source: REDS 1998-99, Arnold, Choe, and Roy (1998)

Table 13: **Bivariate probit estimation**

Dependent variable: Reported pregnancy and residence			
	I: Within family test	**II: Land ownership test**	
		A: Landless head	**B: Land owning head**
Number of own sons	**- 0.333** ***	**- 0.365** **	**- 0.331** ***
	(0.033)	(0.155)	(0.034)
Number of own daughters	**- 0.113** ***	- 0.139	**-0.114** ***
	(0.024)	(0.128)	(0.024)
Other claimants' sons	**0.038** ***	0.083	**0.038** ***
	(0.013)	(0.067)	(0.013)
Other claimants' daughters	- 0.016	- 0.022	- 0.015
	(0.016)	(0.075)	(0.016)
$\beta_{2m} - \beta_{2f}$	**0.054 ****	0.105	**0.053 ***
F statistic	4.15	0.72	3.74

Notes: Values in parentheses are standard errors. Standard errors are clustered at joint family level. *** $p < 1\%$. ** $p < 5\%$. * $p < 10\%$. sN = 7,522 in 599 joint families. Source: REDS 1998-99.