Where there is a will: Fertility behavior and sex bias in large families

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Abstract

Boys and girls in India experience large differences in survival and health outcomes. For example, the 2001 Census reports that the sex ratio for children under six years of age is 927 girls per thousand boys, an outcome that has been attributed to differences in parents' behavior towards their sons and daughters. Most studies rely primarily on cultural factors or biases in economic returns to explain these differences. In this paper, I propose an explanation where bequest motives drive fertility behavior that generates sex-based differences in outcomes even when parents do not explicitly prefer boys over girls. In India's patrilocal rural society, women do not inherit property and heads of joint families aim to retain assets within the family lineage for future generations. I hypothesize that this leads heads to bequeath more land to claimants with more sons, in turn generating a race for sons among adult brothers seeking to maximize their inheritance of agricultural land. I confirm this theoretical prediction using panel data from rural households in India. This strategic fertility behavior implies that girls have systematically more siblings compared to boys, and hence receive smaller shares of household resources, offering an explanation for sex-based differences in 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

To my brother belong your green fields

O father, while I am banished afar...

– Hindi folksong

Boys and girls in India witness 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 largest differences in survival outcomes in the developing world. Among surviving children, boys are more likely than girls to receive immunizations, medical attention when sick, and adequate nutrition (Pande 2003). An extensive literature has addressed these persistent differences between boys and girls, identifying various mechanisms such as differential labor market returns (Rosenzweig and Schultz 1982) and preferences due to culture or tradition (Sen 1990). However, in many such explanations, the causal mechanisms are not clear. In this paper, I show that bequests and associated fertility behavior in an agricultural society can be a significant driver of differential health and survival outcomes for boys and girls, even when parents do not treat daughters and sons differently.

Much of the existing literature suggests that economic or cultural considerations lead to discriminatory behavior by parents. 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 can leave unwanted girls for either future adoption or rearing in state-run orphanages.

We know that existing explanations are incomplete, and the policy responses inadequate, 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). In addition, explanations that rely solely on discrimination in the labor market or in cultural practices fail many tests. Census and National Sample Survey data shows the sex ratio is worse in Indian states where land forms a large part of family assets (Figure 1) and where income from agriculture is high (Figure 2). If economic considerations drive discriminatory behavior, why are outcomes for girls relatively worse in prosperous regions? Also, household level data investigated in this paper indicates that the sex ratio is also worse in large "joint" families, which predominate in rural farming communities. Why is this so, when larger families would arguably provide greater economic security compared to independent families? And what is the role of land ownership in determining the sex ratio?

I address these questions in a model of bequest and fertility behavior among rural, land-owning families in a patrilocal society. Almost universally in India, adult daughters leave their natal family at the time of marriage to join their husband's family and do not inherit land from their fathers. The joint family head divides the land bequest among 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. 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 makes land bequest decisions after observing the number of sons that claimants have, since bequeathing land to a claimant with many daughters and few sons increases the probability that land will eventually leave the family. The 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. Even when indifferent between boys and girls, each claimant will have more children when the other claimants have more boys, 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 same and they do not discriminate between their sons and daughters.

This novel hypothesis for the origin of gender differences in India relies on the nature of norms and institutions associated with land management, marriage and fertility choices. The bequest and fertility behavior as well as the demographic implications of the hypothesis are tested 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. Specifically, I demonstrate 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 % per year. The results are robust to two tests. First, this fertility response is significant if the joint family head owns land, but not otherwise, suggesting that land bequests are motivating strategic fertility. Second, other claimants have an impact on fertility while the head is alive and the land has not yet been distributed, but not after the head's death.

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. 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 (see Rosenzweig and Wolpin (1985), Foster and Rosenzweig (2002), Joshi and Sinha (2003), Edlund and Rahman (2005)).

This paper also adds to the literature on strategic bequest behavior inaugurated by Bernheim, Schleifer, 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, Schleifer, 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.

Strategic fertility behavior poses several challenges for policy-makers aiming to alleviate poorer outcomes for girls in developing countries. The institution of joint families and associated practices remain entrenched in rural society despite efforts to withdraw legal recognition to such family structures. Also, unlike overt acts of sex-selective foeticide and infanticide, individual instances of bequest-motivated differential fertility stopping behavior are arguably difficult to detect or prevent.

This paper is organized as follows. The next section describes the social context of sex discrimination among agricultural households in rural India. Section 3 develops a theoretical model of bequest and fertility behavior in joint families and proposes testable hypotheses. Sections 4 and 5 describe the data, econometric tests and results. Section 6 concludes with discussion of the results.

2 Social context

2.1 Three discrimination puzzles

The reasons for sex differentials in child health and mortality in developing countries remain a puzzle. In their seminal contribution, Rosenzweig and Schultz (1982) propose that sex bias is a rational response to differences in economic returns to men and women. These wage 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).

Misogynistic social and cultural beliefs may also drive male preference (Sen 1990). Gangadharan and Maitra (2003) examine sex bias among different racial and ethnic groups in South Africa and find that sex bias is stronger in the Indian community than in any other group, perhaps due to religious beliefs that privilege men over women.

In both these frameworks, the authors argue that parents actively discriminate in favor of boys and against girls through sex-selective foeticide and infanticide, as well as differences in provision of food and healthcare. Media and popular opinion reinforces this perspective (Dugger 2001; Katz 2006). However, recent analysis has challenged this view. Demographic analysis using the National Family Health Survey 1992 revealed that sex selective foeticide or infanticide cannot be the dominant factor explaining the skewed sex ratio (Bhargava 2003). Most excess male deaths take place during birth or soon after[1], whereas most excess female deaths take place between 7 and 36 months, even after accounting for severe underreporting and misreporting of foeticides and infanticides. Furthermore, analyses of sex-selective abortion, such as Arnold, Kishor, and Roy (2002) and Bhat and Zavier (2007), at best estimate 100,000 such abortions per year, which is insufficient in explaining a gender survival gap of tens of millions. Hence, neglect of infant girls seems to be the main driver of the differences in health and survival outcomes.

The evidence is mixed on whether such neglect represents willful or inadvertent discrimination by parents. Part of the literature argues that parents actively discriminate against daughters in allocating nutrition and health resource (see Das Gupta 1987 and the extensive literature cited in Miller 1981). However, tests of intrahousehold allocation fail to reveal significant bias in behavior. Griffiths, Matthews, and Hinde (2002) reject significant within-family differences in weight by gender.[2] Instead, recent studies present evidence that son-preference manifests itself predominantly in fertility behavior so that the resulting family structure is unfavorable to girls (Basu 1989; Arnold, Choe, and Roy 1998). This fertility behavior takes the form of "stopping rules" where parents have children until they have a certain number of boys are born (Yamaguchi 1989; Clark 2000). Under such rules, the average girl will have systematically more siblings than the average boy, leading to fewer resources and poorer outcomes even with equitable parent behavior. The evidence suggests that stopping rules have significant impact on differential outcomes for girls compared to boys (Basu and de Jong 2008). The origin of these stopping rules is not sufficiently addressed by the literature and is the first puzzle that I will address in this paper.

[1]This is consistent with the medical evidence that the male foetus is much more vulnerable than a female foetus (Gloster and Williams 1992, Andersson and Bergstrom 1998, Andersen et al. 2002).

[2]Also see Jensen 2003 for a list of more such studies.

Rosenzweig and Schultz (1982) argue that discrimination against girls is driven by the economic or social marketplace, which would suggest that the worst outcomes are observed in the most destitute families where the marginal value of an additional son is greatest. However, 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) and more recently Chakraborty and Kim (2008) 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. These contradictory findings constitute the second puzzle addressed in this paper.

Girls experience worse outcomes in large, multi-generational families known as joint families. 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, especially since recent research has shown that children in joint families benefit from higher levels of public good provision (Edlund and Rahman 2005). A possible explanation is that co-resident grandparents transmit traditional ideas on gender roles. George (1997) suggests an active role for the paternal grandmother in performing infanticide. But exactly what these traditional ideas are, why grandparents would believe them, or what motivates grandmothers to perform such gruesome acts is left unanswered. In this paper, the open question of why girls' outcomes are comparatively worse in joint families constitutes the third puzzle.

Thus, explanations for sex discrimination paint at best an incomplete picture, with many assumptions that do not incorporate the nuances of different family structures and social practices in India. This paper offers a new explanation for gender discrimination that specifically address the three puzzles outlined above.

2.2 Rural family structure

The family is the central unit of social organization, production and consumption in most agrarian societies. Much of the development literature treats the "family" as synonymous with the "household" and takes the unitary household as the basis for analysis (see Deaton 1997 for a summary). However, recent surveys that track household formation and dissolution allow researchers to explore more complicated family structures in developing societies.

Caldwell's (1984) basic 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 this paper, I use "independent family" instead). 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 children, along with their wives and minor children (In this paper, I use "joint family" as a shorthand for a joint-stem family). The nuclear family has been the dominant type of family organization in much of the world, except China and North India where joint families are widely observed (Das Gupta 1999).

A widespread social practice in India is that women leave their natal household at the time of marriage and move to their husband's home. As a result of such "patrilocality" or "virilocality", women are considered members only of their family of marriage. Consequently, they have no inheritance rights in their parents' family, neither in law nor in practice since any land given to them would be lost to the family lineage (Agarwal 1998; Mearns 1999; Singh 2005).

Botticini and Siow (2003) show that in patrilocal societies, the family head prefers to leave a bequest of illiquid land only to his sons. Adult sons remain at home throughout their lives and work on the family's land. If assets are distributed to all children at the time of the head's death, sons and daughters have different incentives to exert effort on farm production. The non-resident daughters' effort on the parents' farm is not observable and they might shirk. Sons would not obtain complete rewards from their effort resulting in a free-riding problem with daughters benefiting disproportionately compared to their effort. Hence, they argue, the potential for free-riding explains why daughters receive their share of the bequest as dowries in the form of liquid assets at the time of marriage, rather than in the form of fixed productive assets at the time of the head's death. Chen (2000) offers empirical confirmation for Botticini and Siow's (2003) hypothesis, reporting that only 13% of daughters inherited land after the death of their land-owning fathers.

The farm-based joint family is of particular interest since presumably the demands of agricultural production gave rise to such a structure. Rosenzweig and Wolpin (1985) develop a model where older family members learn how to farm a specific piece of land and transfer this knowledge to their children. This means that using family labor is relatively profitable compared to hired labor. Older and younger family members enter into an implicit contract where the elderly transfer land-specific knowledge to younger family members in return for co-residence. Thus, the model explains both the formation of stem and joint families, as well as the paucity of land sales to non-family members. Land sales occur only in case of extreme distress, particularly weather shocks.[3] Rosenzweig and Wolpin (1985) report that sales of agricultural land are rare in rural India – only 1.75% of all families and 0.39% of stem and joint families in their sample reported any land sales in a year.[4]

Sharing in the consumption of a household public good as well as the possibility of receiving a share in the bequest upon the head's death keeps the sons from splitting away to form their own households. Land is the dominant form of bequest; indeed Das Gupta (1999) reports that the raison d'etre of joint family households is to ensure the continuity of the estate.

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? Various studies propose answers to this question. 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 produce. 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. Additionally, farmers who cultivate their own land do not face classic agency problems and are motivated to exert maximum effort into production (Banerjee, Gertler, and Ghatak 2002).

[3]This explanation for household division is confirmed by Foster and Rosenzweig (2002), who present a model of a farm-based joint family that examines closely the role of public goods as incentives for claimants to remain within the family. When economic distress due to weather shocks or other family-wide factors reduces the provision of household public goods, more couples leave the joint family's household to set up separate households. In addition, household division increases due to claimant inequality in birth order, schooling and the number of sons, but not the number of daughters.

[4]Also see Deininger, Jin, and Nagarajan (2007) for more on rural land market participation in India.

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.

Thus, 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. Hence, equal division is neither the norm nor the law, and adult sons have incentive to alter their behavior to get larger shares of land. I use these features of family behavior to explain the three sex discrimination puzzles presented in section 2.1 – why are gender differences larger in land-owning families, why is the sex ratio worse in joint families, and how do stopping rules arise in fertility behavior?

3 Theory

In Bernheim, Schleifer, and Summers (1985), parents use bequests to induce children to bring their behavior in line with the parents' preferences. The formulation here 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 the parents. In this section, I examine what these two assumptions imply about the household head's bequest and children's fertility behavior. I illustrate how fertility behavior leads to systematic differences in the types of households that girls and boys live in, and how this explains the sex discrimination puzzle. The modeling exercise yields theoretical predictions that can be tested in the data.[5]

I interpret the result from Botticini and Siow (2003) as an explanation for why the head prefers to bequeath land to claimants with more sons in order to perpetuate land ownership within the same lineage. If the head bequeaths any 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.

As an illustration, consider the case of a head who has to choose between two claimants, the first with a boy and a girl and the second with two boys. If the grandsons further have two children each after the head dies, then the probability that the first claimant has at least one grandson and land remains within the family is $3/4$, whereas the probability that the second claimant has at least one grandson and land remains within the family is $15/16$. The head derives additional utility from bequeathing a share of his assets to each claimant, but realizes declining marginal gains from doing so. Suppose $u(\kappa) = \pi\sqrt{\kappa}$, where π is the probability that a claimant has at least one grandson and κ is the fraction of the land bequeathed to the first claimant. Then he maximizes the expected utility from bequests by solving the following problem.

$$\max_{\kappa} \frac{3}{4}\sqrt{\kappa} + \frac{15}{16}\sqrt{1-\kappa} \text{ such that } 0 \leq \kappa \leq 1 \tag{1}$$

The solution to the head's problem is $\kappa = 16/41$ and $1 - \kappa = 25/41$ and the claimant with two sons receives a larger share of the bequest.

[5]The model presented in this section illustrates the essential mechanism of bequest, public good consumption and fertility behavior. To estimate the structural parameters, a joint family model would also incorporate farm production, labor supply, consumption, savings, marriage and residence decisions. The empirical tests in this paper show that residence decisions do not significantly affect bequest or fertility behavior. Modeling and testing other aspects of household decisions await panel datasets that comprehensively measure individual consumption within the family, along with other decisions.

3.1 Model of fertility choice

This section presents a formal model of bequest with endogenous fertility behavior in joint families. The objective of the modelling exercise is to develop a mechanism that links land bequests with fertility behavior, and its influence on health and survival outcomes for girls. The theoretical model generates clear predictions that will be tested empirically in subsequent sections.

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. Heads then observe the claimants' family structure and fertility decisions and make bequest and consumption allocation decisions that maximize their objective function. 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:

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

$$\text{where } \kappa_i \geq 0 \text{ , } \mu_i \geq 0 \text{ and } z \geq 0 \text{ for all } i \tag{2}$$

$$\text{and } \sum_i \kappa_i = 1 \text{ , } \sum_i \mu_i = 1 \text{ , } z + \sum_i \mu_i x_i = I \tag{3}$$

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 react 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 $u_H(\mathbf{.})$ from giving to each claimant. Therefore, the head's problem can be written succinctly as:

$$\max_{\kappa,\mu,z} U_H = \sum_i u_H(\pi_i, \kappa_i, \mu_i, z) \tag{4}$$

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

$$\frac{\partial \pi_i}{\partial m_i} > 0 \text{ and } \pi_i(m_i^0) > 0 \text{ for all } m_i^0 \tag{5}$$

$\{m_i, f_i\}$ 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 (2). Solving the problem for all claimants yields the following reaction functions.

$$\kappa_i = \kappa(\mathbf{m}) \tag{6}$$

$$\mu_i = \mu(\mathbf{m}) \tag{7}$$

$$z = z(\mathbf{m}) \tag{8}$$

The head's preference for bequeathing larger shares of land to claimants with more sons can be written as,

$$\frac{\partial \kappa_i}{\partial m_i} \geq 0, \frac{\partial \kappa_i}{\partial m_{-i}} \leq 0 \tag{9}$$

I term the comparative static in (9) as "strategic bequests". I test for this relationship in the data, which if confirmed, provides the motivation for claimants to have more sons than their brothers.[6]

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

$$\max_{\phi_i} EU_i(x_i, x_i^\delta, \mathbf{n}_i) \tag{10}$$

where expectations are taken over the probability that the head survives in the current period.

[6]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 (9). 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.

$x_i = x_i(\mathbf{n}, \mu, z)$ 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 and public resources (μ, 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 $(\mathbf{n}_i)$.[7]

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

$$m_i = m_i^0 + \mathbf{I}\{\tilde{y} < p\}\phi_i + \tilde{\epsilon}_i^{\phi, m^0} \tag{11}$$

$$f_i = f_i^0 + \mathbf{I}\{\tilde{y} > p\}\phi_i + \tilde{\epsilon}_i^{\phi, f^0} \tag{12}$$

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 $\phi_i = 1$. $\tilde{\epsilon}_i^{\phi, m^0} \in \mathbf{I}$ is a discrete fertility shock whose distribution depends on ϕ_i and m_i^0. Similarly, the distribution of $\tilde{\epsilon}_i^{\phi, f^0} \in \mathbf{I}$ 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.

$$\phi_i^* = \phi_i(\mathbf{m}^0) \tag{13}$$

$$\phi_{-i}^* = \phi_{-i}(\mathbf{m}^0) \tag{14}$$

[7]The claimant can draw utility from current consumption while foreseeing his future role as a household head if his comprehensive utility consists of two separable parts - utility from consumption as a claimant and utility from bequests as a head.

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

3.2 Impact of fertility on family structure

Strategic bequests that lead to more pregnancies do not by themselves imply unequal sex outcomes. This section shows the demographic implications of strategic bequests over time 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 some standard assumptions on the form of the head's and claimants' utility functions. I 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 (13) and (14).

$$\frac{\partial U_H}{\partial \kappa_i} \geq 0, \frac{\partial^2 U_H}{\partial \kappa_i^2} < 0, \frac{\partial U_i}{\partial x_i} \geq 0, \frac{\partial^2 U_i}{\partial x_i^2} < 0 \tag{15}$$

where x represents the claimant's consumption of household goods as well as children. I further assume that there exists an $\hat{m}_i$ 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.

$$Pr\{\phi_i = 1 \mid m_i^0, m_{-i}^0\} > Pr\{\phi_i = 1 \mid m_i^0 + 1, m_{-i}^0\} \tag{16}$$

$$Pr\{\phi_i = 1 \mid m_i^0, m_{-i}^0 + 1\} > Pr\{\phi_i = 1 \mid m_i^0, m_{-i}^0\} \tag{17}$$

I term the theoretical prediction in equation (17) "strategic fertility". Given this result, 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 (16) and (17) imply that B, who has a new daughter, has greater incentive than A to have another child.

$$Pr\{\phi_B = 1 \mid m_A, m_B\} > Pr\{\phi_A = 1 \mid m_A, m_B\} \tag{18}$$

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

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

$$\text{No. of siblings for average boy} = \frac{\frac{3}{2}Pr\{\phi_A = 1\} + \frac{1}{2}Pr\{\phi_B = 1\}}{1 + \frac{1}{2}Pr\{\phi_A = 1\} + \frac{1}{2}Pr\{\phi_B = 1\}} \tag{20}$$

$$\frac{\text{No. of siblings for average girl}}{\text{No. of siblings for average boy}} = \frac{\frac{1}{2}Pr\{\phi_A = 1\} + \frac{3}{2}Pr\{\phi_B = 1\}}{\frac{3}{2}Pr\{\phi_A = 1\} + \frac{1}{2}Pr\{\phi_B = 1\}} > 1 \tag{21}$$

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.

$$E(\text{Number of siblings for average girl}) > E(\text{Number of siblings for average boy}) \tag{22}$$

As a result, the average girl will have systematically more siblings than the average boy to share her resources. This means that 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.

4 Rural Economic and Demographic Survey

Testing the theoretical predictions presented in Section 3 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 9 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 8 villages that were located in Jammu and Kashmir, where a violent separatist movement perhaps made the survey difficult.[8]

I use data from the 1998-99 wave to test the theory presented in Section 3. Previous waves are used to categorize households as 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 household heads 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.

[8]Since the separatist movements in Assam or Jammu and Kashmir are unlikely to be related to family dynamics, I am not concerned about the missing villages as 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.

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 that household is independent or part of a larger joint family. The heads also provided detailed information on the source, value and extent of their 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.

In both the head and women's survey, since respondents report dates associated with events such as births, deaths and household division, I recover an annual retrospective panel dataset from a single wave of observations in 1998-99. An important feature of the dataset is a detailed fertility history for each woman that records whether or not the claimant reported a pregnancy in every period and the number of living children in that period. 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 (equation 9). 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 (equation 17).

Figure 3 shows three generations of a joint family. The bequest dataset contains the first generation as the head, and the second generation as claimants. The fertility dataset is constructed using the second generation as the head, and the third generation as 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.

Specific features of these two datasets are described in the next section that tests for strategic bequest and strategic fertility behavior. I then analyze whether the number of excess siblings for girls compared to boys in joint families exceeds the number in stem families.

5 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 econometric exercise has three objectives. The first objective is to confirm the strategic bequest motive of equation (9), particularly whether the claimant's share of a bequest is influenced positively by the number of sons. This establishes the value of sons to claimants in the bequest game. Strategic allocations of household public and private consumption goods are not tested since these are not observed in the REDS data. The second objective is to test strategic fertility behavior predicted in equation (17), 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. The third objective is to 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. Note that I do not estimate the impact of strategic fertility on actual survival or health outcomes.

5.1 Strategic bequests

5.1.1 Data: Strategic bequests

The bequest dataset contains cross-sectional information as reported at the time of the head's death. It consists of those land-owning households that were part of a single land-owning unit in 1981-82, but had split into at least two households by 1998 following the head's death in the interim.[9] 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.

[9]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.

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.[10] Note that the average number of sons per claimant is 1.1, which is significantly greater than the average number of daughters (0.9).

5.1.2 Specification: Strategic bequests

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 ($\sum_{k \neq i} \mathbf{n_{kj}}$). The bequest share κ_{ij} is censored below 0 and above 1. Therefore, I specify the following dual-censored tobit model.

$$\kappa_{ij} = \begin{cases} 0 & \text{if } \kappa^*_{ij} \leq 0 \\ \kappa^*_{ij} & \text{if } 0 < \kappa^*_{ij} < 1 \\ 1 & \text{if } 1 \leq \kappa^*_{ij} \end{cases} \tag{23}$$

Here, κ^*_{ij} is a latent variable such that

$$\kappa^*_{ij} = \alpha_0 + \alpha_1 \mathbf{n_{ij}} + \alpha_2 \sum_{k \neq i} \mathbf{n_{kj}} + \alpha_3 \mathbf{n_{ij}} * r_{ij} + \alpha_4 \sum_{k \neq i} \mathbf{n_{kj}} * r_{ij} + \alpha_5 \mathbf{X_{ij}} + \alpha_6 \mathbf{Y_j} + \xi_{ij} \tag{24}$$

where $\mathbf{n}_i = [m_i, f_i]'$, $\alpha_1 = [\alpha_{1m}\ \alpha_{1f}]$, $\alpha_2 = [\alpha_{2m}\ \alpha_{2f}]$, $\alpha_3 = [\alpha_{3m}\ \alpha_{3f}]$ and $\alpha_4 = [\alpha_{4m}\ \alpha_{4f}]$

To confirm the strategic bequest hypothesis, I expect that the bequest share rises in the number of own sons, $\alpha_{1m} > 0$, and falls in the number of other claimants' sons, $\alpha_{2m} < 0$, corresponding to the theoretical predictions in equation (9). 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. The impact of moving away is theoretically

[10] These land holdings are consistent with the national average holding of 1.67 hectares in 1981-82, 1.34 hectares in 1991-92 and 1.06 hectares in 2002-03 reported in Govt. of India (2006)

ambiguous because splitting from the head's household might indicate 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.

This specification must be qualified by controlling for the claimant's residence choice r_{ij} and other observed claimant-specific $\mathbf{X_{ij}}$ factors that might impact bequest preferences. $\mathbf{X_{ij}}$ consists of claimant-specific characteristics such as age at time of inheritance, education and wife's education. 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. $\mathbf{Y_j}$ includes family specific factors such as the head's education and other demographic characteristics. A dummy variable indicates whether the head was a farmer. 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 normally with zero mean.

One limitation of this specification is that dependent variables are not independent across observations. Specifically, in a sample of all claimants, the bequest of one claimant is simply the residual share from the other claimants. Therefore, one approach I take is to separately estimate equation (23) for first-born and other claimants since shares will not be correlated in a sample that includes only first-born claimants.

However, that approach does not use the information that claimant shares must add to one. Therefore, another approach I try is to 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 the share of land bequeathed to one of the claimants. Due to small number of observations, I cannot estimate a similar model for families with more than two claimants, or with more categories.

5.1.3 Results: Strategic bequests

The results from specification (23), where I test for the influence of family structure on received bequest shares, are presented in Table 4. Coefficients from a Tobit estimation cannot be directly interpreted as percentages. However, since the number of censored observations is small, I expect that Tobit is a close approximation of OLS, and therefore interpret the coefficients in Table 4 for now as percentages. Column I reports the results for all claimants within the household, and shows that an additional son increases the claimant's share of the land bequest by 1%. This mirrors the cumulative increase in bequest share for the other claimants when they have an additional son (1.2%). 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. Column I also shows that a claimant's birth order has a large influence on the bequest share received by a claimant. A claimant increases his bequest share by 8% with an improvement of one position in the birth order.[11]

Note that the claimant's residence away from the head's household 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 concurrently in order to receive a larger inheritance. From an econometric perspective, this suggests that the results from a probit estimation of fertility choice in equation (5.2.2) should be similar to the joint estimation of fertility and residence choice (Appendix B).

Columns II and III repeat the estimation with subsamples of claimants who are first and higher in the birth order, respectively. The coefficients for α_{1m} and α_{4m} are both positive and significant, indicating that an additional son adds approximately 2% to a claimant's bequest. Interestingly, heads seem to value the daughters of claimants with high birth order, although not as much as sons.

The dataset does not report the value of land inherited. However, data on the value of land purchased during the reporting period is available. I calculate the value of the bequest distributed by the household head as Rs. 2.4 million, measured in 1999.[12] Thus, a 2% increase in bequest share suggests that the value of an additional son is Rs. 47,000 to the average claimant in the bequest dataset.

[11] 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.

[12] Rs. 43.19 = US$ 1 on 07/07/08.

Table 5 reports the results of the multinomial logit estimation for two-claimant joint families. Again, this specification incorporates the fact that bequest shares across claimants sum to one. Additional sons help a claimant receive a greater share of land although the impact is small. An additional son can help break an equal division tie so that a claimant receives between 0.5 and 0.55% of the bequest share, but does not help a claimant to receive larger shares of the bequest. While these results support the strategic bequest hypothesis and are in line with the estimates presented in the tobit specification, the small number of observations in each category implies that they should be interpreted with caution.

These results establish that the number of own and other claimants sons are important factors determining the bequest received by the claimant. 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.

5.2 Strategic fertility

To test the strategic fertility hypothesis, I propose three tests. First, "within-family fertility" tests the strategic fertility hypothesis directly for all claimants and families. Second, "land ownership" tests the impact of land ownership by the head on strategic fertility. Finally, "head's death" tests the impact of the head's death on strategic fertility.

5.2.1 Data: Strategic fertility

Each observation in the fertility dataset consists of a man who is older than 15 years of age. Each 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 an independent household.

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.[13] Therefore, the sample is restricted to

[13]Recalled fertility data suffers from bias from two main sources (Schultz 1972). The primary reason is that

the 1992-98 time period which leaves 43,612 claimant-family-year observations in the panel from 5,090 families over seven years.

Claimants might live within the household occupied by the joint family head or set up an independent household. $r_i \in \{0, 1\}$ represents the claimant's residence within or outside the head's household respectively. In the survey, multiple household heads who originated from a single household in the 1981-82 wave might either be independent family heads or claimants in a joint family. This status is based on the circumstances of departure. Fortunately, the REDS dataset collects information on household division. 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. Consistent with observed bequest behavior, split off sons retain status as claimants in their father's household.

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, 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. Both these events are assumed to 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

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.

other family types.[14].

5.2.2 Within-family fertility

Specification

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 (17). The other claimants' daughters are neither future heirs in the family lineage, nor direct economic costs or benefits to the claimant. Hence, they ought not to have a large or significant impact on claimant's own fertility. To test these two propositions, I specify a probit model with a binary outcome θ that is 1 if a claimant i in joint family j reports a pregnancy in year t.

$$\theta_{ijt} = \begin{cases} 0 & \text{if } \theta^*_{ijt} \leq 0 \\ 1 & \text{if } 0 < \theta^*_{ijt} \end{cases} \tag{25}$$

Here, θ^*_{ijt} is a latent variable such that

$$\theta^*_{ijt} = \beta_0 + \beta_1 \mathbf{n}_{ijt} + \beta_2 \sum_{k \neq i} \mathbf{n}_{kjt} + \beta_3 \mathbf{X}_{ijt} + \beta_4 \mathbf{V}_{ij} + \beta_5 r_{ijt} * \mathbf{n}_{ijt} + \beta_6 r_{ijt} * \sum_{k \neq i} \mathbf{n}_{kjt} + year_t + \mu_j + \epsilon_{ijt} \tag{26}$$

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 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, equation (16) predicts that $\beta_{1m} < \beta_{1f}$. Strategic fertility is identified by the components of β_2. In particular, equation (17) predicts that $\beta_{2m} > 0$ and is statistically significant, but β_{2f} is close to zero and not significant. The coefficient β_5 indicates the impact of the claimant's own sons and daughters if he is living in a split-off household. β_6 indicates the impact of the other claimants sons and daughters when the claimant is split off.

[14]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).

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 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).

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 check for this, I exploit the panel aspect of the dataset and also specify a probit random effects model where

$$\theta^*_{ijt} = \beta_0+\beta_1\mathbf{n}_{ijt}+\beta_2\sum_{k\neq i}\mathbf{n}_{kjt}+\beta_3\mathbf{X}_{ijt}+\beta_4\mathbf{V}_{ij}+\beta_5 r_{ijt}*\mathbf{n}_{ijt}+\beta_6 r_{ijt}*\sum_{k\neq i}\mathbf{n}_{kjt}+family_{jt}+year_t+\epsilon_{ijt} \quad (27)$$

In 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 (25).

Results

Table 6 presents the results of the within-family test of strategic fertility specified in section 5.2.2. In the within-family test, the strategic fertility model predicts that the number of other claimants' boys has a positive and significant impact on the likelihood of reporting a pregnancy, i.e. $\beta_{2m} > 0$. Simultaneously, the other claimants' girls should have a small and statistically insignificant impact on the claimant's fertility, i.e. β_{2f} is small. Column I reports marginal-effects probit estimates from the specification in equation (25). 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% with an additional son, and by 2.2% with an additional daughter. In contrast, an additional son for the other claimants increases the probability of a pregnancy by 0.85% in a year, a result that is significant at the 1% level. The other claimants' daughters have a small impact on the claimant's fertility (0.5%) that is statistically indistinguishable from the null. This result establishes the basic validity of the strategic fertility hypothesis.

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% for all claimants to 4.6% 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%) of splitting on the marginal fertility impact of the other claimants' sons.

Column II in Table 6 reports the results on the random effects probit model specified in equation (27). This model includes controls for 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 results from this model are not much 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.[15] In addition, the other claimants' daughters seem to be statistically different from the null at the 10% level. However, since the associated point estimate is smaller than the impact of either own children or the other claimants' sons, the validity of the strategic fertility hypothesis is maintained.

Table 7 presents the estimates for the marginal effect on own fertility while fixing the other claimants' sons. The first row of coefficients indicates that the value of an additional son for a claimant is larger when the other claimants have more sons, than when the other claimants have fewer 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.

[15] Note that the coefficients in Column II are different from Column I since this column reports probit, not marginal effects probit results.

The third row of coefficients in Table 7 confirms the earlier result that a claimant is more likely to report a pregnancy when the other claimants have more sons. I cannot conclude that this result is driven by any particular number of other claimants' sons although the effect is greater when the other claimants have more sons confirming the theoretical prediction in the previous paragraph.

5.2.3 Land ownership

Specification

This test uses Bernheim, Shleifer and Summer's (1985) prediction that the strategic bequest game is impacted by the size of the bequest. Correspondingly, 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. Indexing $l \in [0, 1]$, I estimate the following probit model where the model in equations (25) and (26) is interacted with land ownership.

$$\theta_{ijt} = \begin{cases} 0 & \text{if } \theta^*_{ijt} \leq 0 \\ 1 & \text{if } 0 < \theta^*_{ijt} \end{cases} \tag{28}$$

Here, θ^*_{ijt} is a latent variable such that

$$\begin{aligned} \theta^*_{ijt} &= \sum_l \gamma^l_0 * I^l + \sum_l \gamma^l_1 (I^l * \mathbf{n}_{ijt}) + \sum_l \gamma^l_2 (I^l * \sum_{k \neq i} \mathbf{n}_{kjt}) + \sum_l \gamma^l_3 (I^l * \mathbf{X}_{ijt}) + \sum_l \gamma^l_4 (I^l * \mathbf{V}_{ij}) \\ &+ \sum_l \gamma^l_5 (I^l * r_{ijt} * \mathbf{n}_{ijt}) + \sum_l \gamma^l_6 (I^l * r_{ijt} * \sum_{k \neq i} \mathbf{n}_{kjt}) + \sum_l (I^l * year_t) + \epsilon_{ijt} \end{aligned} \tag{29}$$

$$\text{where } \gamma^l_1 = [\gamma^l_{1m}\ \gamma^l_{1f}],\ \gamma^l_2 = [\gamma^l_{2m}\ \gamma^l_{2f}],\ \gamma^l_5 = [\gamma^l_{5m}\ \gamma^l_{5f}],\ \gamma^l_6 = [\gamma^l_{6m}\ \gamma^l_{6f}]$$

I^l is an indicator variable such that $I^0 = 1$ if the head does not own any land and $I^1 = 1$ if the head owns land. The parameters of interest are γ^0_{2m}, γ^0_{2f}, γ^1_{2m} and γ^1_{2f}. If strategic fertility operates only when the head has land that can be bequeathed but not otherwise, then I expect that $\gamma_{2m}^1 > 0$ and significant but γ_{2m}^0 is close to zero and not significant. γ_{2f}^0 and γ_{2f}^1 should both be close to zero and insignificant. As in the previous section, I also estimate a probit model with family random effects and report those results.

Results

Table 8 presents the results of the test of strategic fertility specified in section 5.2.3. Column I reports results from marginal effects probit, and Column II from random effects probit models respectively. In each column, the set of coefficients under 'A' represent claimants in joint families where the head does not own any land. The coefficients under 'B' represent claimants in joint families with a land-owning head.

The basic result corresponding to specification (28) as well as the random effect version is that while the claimant's own family structure is a statistically significant determinant 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% in land owning families. This estimate is close to the 0.85% increase in the fertility rate reported in section 5.2.2. The same coefficient for landless families is larger, i.e., 2.3%. However, it is imprecisely estimated, and cannot be statistically distinguished from zero. In addition, it 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 support the notion that strategic fertility is primarily a phenomenon among land-owning families, offering an explanation why the sex ratio is relatively worse in such families.

5.2.4 Head's death

Specification

The final 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. Assuming that the head's death is not associated with fertility behavior, selection into the sample is random for the purposes of this test. Within the sample, other claimants' sons ought to positively impact a claimant's fertility only while the head is still alive and has not distributed the bequest. Once the head dies and distributes the bequest, claimants have no further incentive for strategic fertility. The following probit specification tests this proposition. This model interacts the specification in equations (25) and (26) with an indicator variable for the head's death.

$$\theta_{ijt} = \begin{cases} 0 & \text{if } \theta^*_{ijt} \leq 0 \\ 1 & \text{if } 0 < \theta^*_{ijt} \end{cases} \tag{30}$$

Here, θ^*_{ijt} is a latent variable such that

$$\begin{aligned} \theta^*_{ijt} &= \sum_d \gamma_0^d I^d + \sum_d \gamma_1^d (I^d * \mathbf{n}_{ijt}) + \sum_d \gamma_2^d (I^d * \sum_{k \neq i} \mathbf{n}_{kjt}) \\ &+ \sum_d \gamma_3^d (I^d * \mathbf{X}_{ijt}) + \sum_d \gamma_4^d (I^d * \mathbf{V}_{ij}) + \sum_d (I^d * year_t) + \mu_j + \epsilon_{ijt} \end{aligned} \tag{31}$$

where $\gamma_1^l = [\gamma_{1m}^l \; \gamma_{1f}^l]$, $\gamma_2^l = [\gamma_{2m}^l \; \gamma_{2f}^l]$, $\gamma_5^l = [\gamma_{5m}^l \; \gamma_{5f}^l]$, $\gamma_6^l = [\gamma_{6m}^l \; \gamma_{6f}^l]$

where $d \in [0, 1]$. $I^0 = 1$ and $I^1 = 0$ before the head dies and $I^0 = 0$ and $I^1 = 1$ afterwards. The parameters of interest are γ_{2m}^0, γ_{2f}^0, γ_{2m}^1 and γ_{2f}^1. If strategic fertility is significant before the head's death but not so afterwards, then I expect that $\gamma_{2m}^0 > 0$ and significant while γ_{2m}^1 is small and not significant. γ_{2f}^0 and γ_{2f}^1 ought to be close to zero and insignificant since the other claimants' daughters are not factors in the claimant's fertility decision.

Results

Section 5.2.4 specifies that a claimant's fertility ought to be dependent on other claimants' family structure only before the head's death and distribution of the bequest. Once the claimant has received his bequest, he will no longer participate in the strategic fertility game. Table 9 reports the results of this test from a marginal effects probit model. The coefficients under 'A' represent the impact of family structure before the head's death, and the coefficients in 'B' represent the impact 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 result is significant only before the head's death.

In addition, the other claimant's sons have a positive and statistically significant impact on own fertility before the head's death. Unusually, the other claimants' daughters have a negative impact on fertility before the head's death, although it is not clear why this is the case. The point estimates imply that an additional own son decreases the probability of reporting a pregnancy by 3.3% before the head's death, but by 13.5% after the head's death. One reason for this large difference is that even controlling for age, the claimant is more likely to have higher order births 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 additional son for the other claimants increases fertility by 1.3% before the head's death, an estimate that is significant at the 5% level, but has virtually no impact following the head's death. This implies that claimants' care about each others' fertility only insofar that the head is alive and has not yet distributed his land, but not so once the head dies and land division is complete.

5.3 Implications of strategic fertility

The results from the previous sections confirm that land bequests in joint families motivate strategic fertility behavior. This behavior implies that a claimant will stop having children sooner when he has many boys rather than when he has many girls. However, as previously discussed in Section 3.2, this result by itself does not guarantee differences in outcomes for girls and boys. For this, I propose that 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 parents' behavior towards children of different gender or in resource allocations, the average girl will receive smaller share of resources than the average boy, explaining poorer outcomes.

To see this in the fertility dataset, I check whether the average girl indeed has more siblings than the average boy. In the following equations, f_{ij} and m_{ij} is the number of sons and daughters 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.

$$\bar{s}_f = \frac{\sum_{i,j}(s_{ij} * f_{ij})}{\sum_{i,j} f_{ij}} \text{ and } \bar{s}_m = \frac{\sum_{i,j}(s_{ij} * m_{ij})}{\sum_{i,j} m_{ij}} \tag{32}$$

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. The number of siblings for the average girl ($\bar{s}_f$) in a joint family is 2.761 whereas the number of siblings for the average boy in a joint family is 2.481. 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. The number of siblings for the average girl in a stem family (2.757) is close to the number of siblings for the average girl in a joint family (2.761). However, the difference in the number of siblings for the average boy in a stem family (2.600) and the number of siblings for the average boy in a joint family (2.481) is large.[16] This is consistent with the theory presented in Section 3 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.

[16] Note that the excess siblings for the average girl is not a trivial outcome of a sex ratio skewed against girls. If girls and boys are randomly assigned to households, then the average girl will have the same number of siblings as the average boy regardless of the sex ratio.

6 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 outcomes more effectively than sex-selective foeticide and infanticide.

I test both the strategic bequest and strategic fertility hypotheses. I confirm that heads prefer 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 also salient before the head's death and distribution of the bequest, compared to families where the inheritance has been received. Although estimating the precise impact of this behavior within joint families on sex differences in health and survival outcomes awaits advances in data collection, these results provide a comprehensive explanation why such differences are greater in joint families than other family structures.

The results should be read with two 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. Moreover, sex bias might be motivated for reasons other than bequests. The impact of strategic bequests and fertility are congruent to these reasons, not in opposition to them.

Second, my 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 so central to the production process, or that have alternative types of social institutions.

The model presented in this paper makes a number of assumptions and simplifications due to constraints in the data as well as practical considerations. The most salient assumptions were to ignore farm production, labor supply, consumption, savings and marriage decisions that are also part of the economics of the joint family household. This has two major implications. First, I cannot comment on the dynamics of fertility behavior in independent households that do not have adult claimants, and hence different sets of labor force participation and consumption decisions than stem and joint family households. Second, I cannot perform simulations that predict the impact of specific policies that seek to redress gender differences in outcomes if those policies are also expected to alter other household decisions.

Relaxing these assumptions requires the development of a full-scale model of intra-family bargaining with forward-looking agents that extends two-agent bargaining models such as those developed by Chiappori (1992), Lundberg and Pollak (1993) and Friedberg and Stern (2007) to a more elaborate family structure. Estimation of all the structural parameters of such a model would be greatly aided by advances in data collection, particularly the allocation of resources within the household. However, elements of the full-scale model can be tested using reduced form techniques and available time use data. For example, the value of a son as a future heir might imply that the birth of a boy increases consumption of leisure for a claimant and his wife. From the perspective of marriage, the status of daughters as residual claimants who inherit land only when the head has no sons implies that women with no brothers would be attractive in the matrimonial market.

However, 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. Since then, these states have been at the forefront of large advances in female survival and health. A recent amendment to the Hindu Succession Act (2005) extended these rights nationally. This ought to increase the bargaining power of daughters in the bequest “game” as they are regarded as claimants in their own right. 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.

References

Agarwal, B. (1998). *A Field of One's Own: Gender and Land Rights in South Asia*. Cambridge UK: Cambridge University Press.

Andersen, C., A. Byskov, M. Fukuda, K. Fukuda, and T. Shimizu (2002). Parental periconceptional smoking and male: Female ratio of newborn infants. *The Lancet 359*(9315), 1407–1408.

Andersson, R. and S. Bergstrom (1998). Is maternal malnutrition associated with a low sex ratio at birth? *Human Biology 70*(6), 1101–1106.

Arnold, F., M. Choe, and T. Roy (1998). Son preference, the family-building process and child mortality in India. *Population Studies 52*, 301–315.

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

Banerjee, A., P. Gertler, and M. Ghatak (2002). Empowerment and efficiency: The economics of agrarian reform. *Journal of Political Economy 110*(2), 239–280.

Basu, A. (1989). Is discrimination in food really necessary for explaining sex differentials in childhood mortality? *Population Studies 43*(2), 193–210.

Basu, D. and R. de Jong (2008). Son preference and gender inequality. Ohio State University Working Paper.

Bernheim, D., A. Schleifer, and L. Summers (1985). The strategic bequest motive. *Journal of Political Economy 93*, 1045–1076.

Bhargava, A. (2003). Family planning, gender differences and infant mortality: Evidence from Uttar Pradesh, India. *Journal of Econometrics 112*(1), 225–240.

Bhat, P. and A. Zavier (2007). Factors influencing the use of prenatal diagnostic techniques and the sex ratio at birth in India. *Economic and Political Weekly 42*(24), 2292–2303.

Botticini, M. and A. Siow (2003). Why dowries? *American Economic Review 93*, 1385–1398.

Caldwell, J., P. Reddy, and P. Caldwell (1984). The determinants of family structure in rural South India. *Journal of Marriage and the Family 46*(1), 215–229.

Census of India (2001). *Provisional Population Totals: India*. Delhi: Office of the Registrar General, India.

Chakraborty, T. and S. Kim (2008). Caste, kinship and sex ratios in India. National Bureau of Economic Research Working Paper No. 13828.

Chen, M. (2000). *Perpetual Mourning: Widowhood in Rural India*. Philadelphia PA: University of Pennsylvania Press.

Chiappori, P. (1992, June). Collective labor supply and welfare. *Journal of Political Economy 100*(3), 437–467.

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

Das Gupta, M. (1987). Selective discrimination against female children in rural Punjab. *Population and Development Review 13*(1), 77–100.

Das Gupta, M. (1999). Lifeboat versus corporate ethic: Social and demographic implications of stem and joint families. *Social Science and Medicine 49*, 173–184.

Deaton, A. (1997). *The Analysis of Household Surveys: A Microeconometric Approach to Development Policy*. Washington DC: World Bank.

Deininger, K., S. Jin, and H. Nagarajan (2007). Determinants and consequences of land sales market participation: Panel evidence from India. World Bank Policy Research Working Paper No. 4323.

Dreze, J. and A. Sen (2002). *India: Development and participation*. New Delhi: Oxford University Press.

Dugger, C. (2001). Abortions in India spurred by sex test skew the ratio against girls. *The New York Times April 22*.

Edlund, L. and A. Rahman (2005). Household structure and child outcomes: Nuclear vs. extended families - Evidence from Bangladesh. Working Paper.

Feder, G. and T. Onchan (1987). Land ownership security and farm investment in Thailand. *American Journal of Agricultural Economics 69*(2), 311–320.

Foster, A. and M. Rosenzweig (2002). Household division and rural economic growth. *Review of Economic Studies 69*, 839–869.

Foster, A. and M. Rosenzweig (2003). Economic growth and the rise of forests. *Quarterly Journal of Economics 118*(2), 601–637.

Friedberg, L. and S. Stern (2007). Marriage, divorce, and asymmetric information. Working paper.

Gangadharan, L. and P. Maitra (2003). Testing for son preference in South Africa. *Journal of African Economies 12*(3), 371–416.

George, S. (1997). Female infanticide in Tamil Nadu, India: From recognition back to denial? *Reproductive Health Matters 5*(10), 124–132.

Gloster, S. and R. Williams (1992). Human sex ratio as it relates to caloric availability. *Social Biology 39*(3-4), 285–291.

Govt. of India (1998). *Household assets and indebtedness of social groups as on 30.6.91, Debt and investment survey, NSS Report No. 432*. New Delhi: National Sample Survey Organisation. Available at `http://mospi.nic.in/mospi_nsso_rept_pubn.htm`.

Govt. of India (2006). *Some aspects of operational land holdings in India, 2002-03, NSS Report No. 492*. New Delhi: National Sample Survey Organisation. Available at `http://mospi.nic.in/mospi_nsso_rept_pubn.htm`.

Griffiths, P., Z. Matthews, and A. Hinde (2002). Gender, family, and the nutritional status of children in three culturally contrasting states of India. *Social Science and Medicine 55*(5), 775–790.

Jensen, R. (2003). Equal treatment, unequal outcomes? Generating gender inequality through fertility behavior. Working Paper. Kennedy School of Government, Harvard University.

Joshi, S. and N. Sinha (2003). The effect of household partition on childrens work and schooling in rural Bangladesh. Working paper.

Katz, N. (2006). Abortion in India: Selecting by gender. *The Washington Post May 20*, B09.

Lundberg, S. and R. Pollak (1993). Separate Spheres and the Marriage Market. *Journal of Political Economy 101*, 988–1010.

Mahajan, A. and A. Tarozzi (2007). Child nutrition and son preference in India during the Nineties. *Economic Development and Cultural Change 55*(3), 441–486.

Mearns, R. (1999). Access to land in rural India. World Bank Working Paper No. 2123.

Miller, B. (1981). *The Endangered Sex: Neglect of Female Children in Rural North India*. Ithaca NY: Cornell University Press.

Pande, R. (2003). Selective gender differences in childhood nutrition and immunization in rural India: The role of siblings. *Demography 40*(3), 395–418.

Rose, E. (1999). Consumption smoothing and excess female mortality in rural India. *Review of Economics and Statistics 81*(1), 41–49.

Rosenzweig, M. and T. Schultz (1982). Market opportunities, genetic endowments, and intrafamily resource distribution: Child survival in rural India. *American Economic Review 72*(4), 803–815.

Rosenzweig, M. and K. Wolpin (1985). Specific experience, household structure, and intergenerational transfers: Farm family land and labor arrangements in developing countries. *Quarterly Journal of Economics 100*(Supl.), 961–987.

Schultz, T. (1972). Retrospective evidence of a decline of fertility and child mortality in Bangladesh. *Demography 9*(3), 415–430.

Sen, A. (1990). More than 100 million women are missing. *The New York Review of Books 37*(20).

Sen, A. and S. Sengupta (1983). Malnutrition of rural children and the sex bias. *Economic and Political Weekly 18*(19), 855–864.

Singh, K. (2005). Amendments to the Hindu Succession Act: Are they enough to bring about gender equality? *Combat Law 3*(5).

Srinivas, M. (1976). *The Remembered Village*. Berkeley CA: University of California Press.

The PROBE Team (1999). *Public report on basic education in India*. New Delhi: Oxford University Press.

Yamaguchi, K. (1989). A formal theory for male-preferring stopping rules of childbearing: Sex differences in birth order and in the number of siblings. *Demography 26*(3), 451–465.

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.

$$\max_{\mu_1,\ldots,\mu_N,z} \sum_i \eta_i(\mathbf{m}) u_i(\mu_i X, z) \tag{33}$$

$$\text{such that } z + \sum_i \mu_i x_i = I \text{ and } \sum_i \mu_i = 1 \tag{34}$$

Then the first order conditions yield

$$\eta_i(\mathbf{m}) \frac{\partial u_i}{\partial \mu_i} = \eta_{-i}(\mathbf{m}) \frac{\partial u_{-i}}{\partial \mu_{-i}} \tag{35}$$

$$\text{or } \frac{\eta_i(\mathbf{m})}{\eta_{-i}(\mathbf{m})} = \frac{MU_{-i}}{MU_i} \tag{36}$$

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

$$\frac{\partial \mu_i}{\partial m_i} \geq 0, \frac{\partial \mu_i}{\partial m_{-i}} \leq 0 \tag{37}$$

B Robustness check for endogenous residence choice

One concern with the tests of strategic fertility presented in section 5.2 is the possibly endogenous determination of residence choice with fertility. As outlined in Section 2.2, 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 5.1.3 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 (25) respectively.

Column I in table 10 reports the results from joint determination of fertility and residence choice in the within family test presented in section 5.2.2. Column II reports the results from joint determination of fertility and residence choice in the land ownership test (section 28), 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 6, confirming that residence choice is not a significant factor in the strategic fertility game.

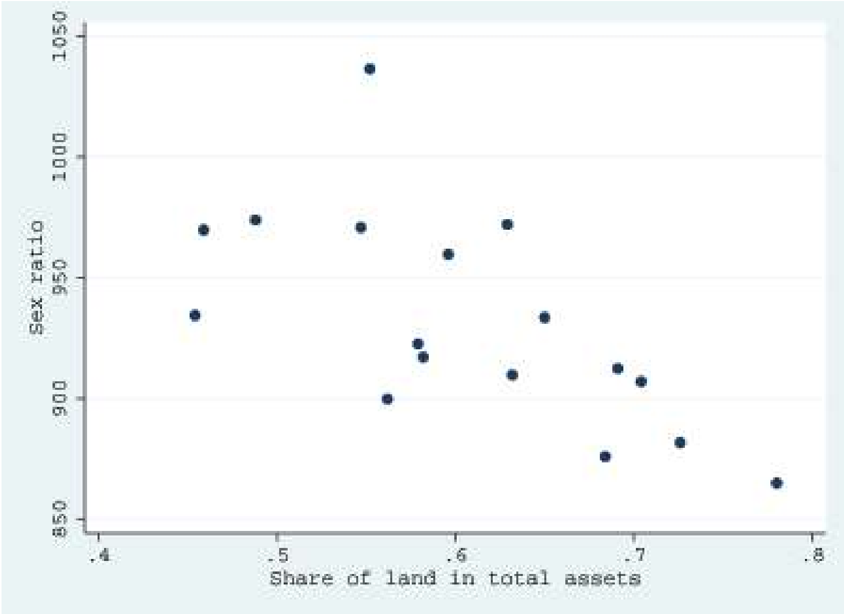

Figure 1: **Importance of land vs. Sex ratio**
Source: Govt. of India (1998) and Census of India (2001).

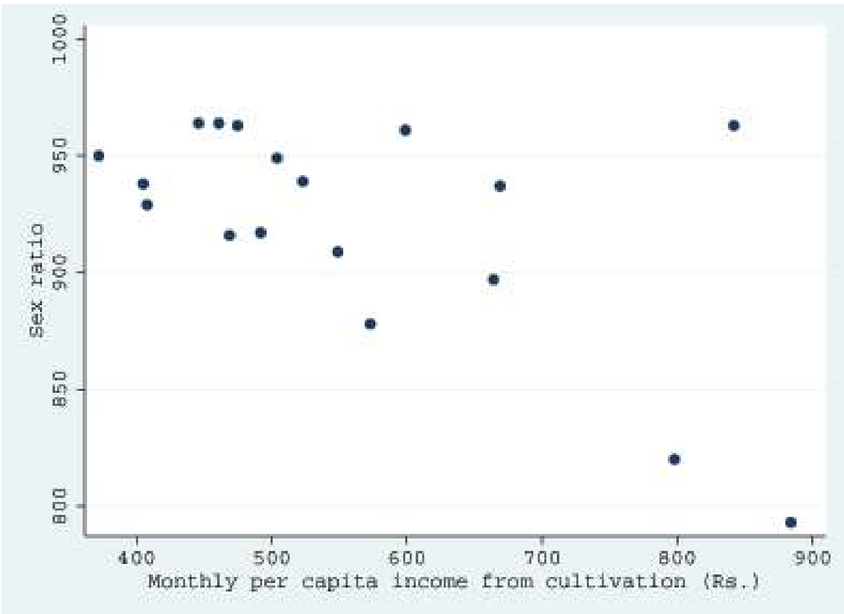

Figure 2: **Agricultural income vs. Sex ratio**
Source: Govt. of India (1998) and Census of India (2001).

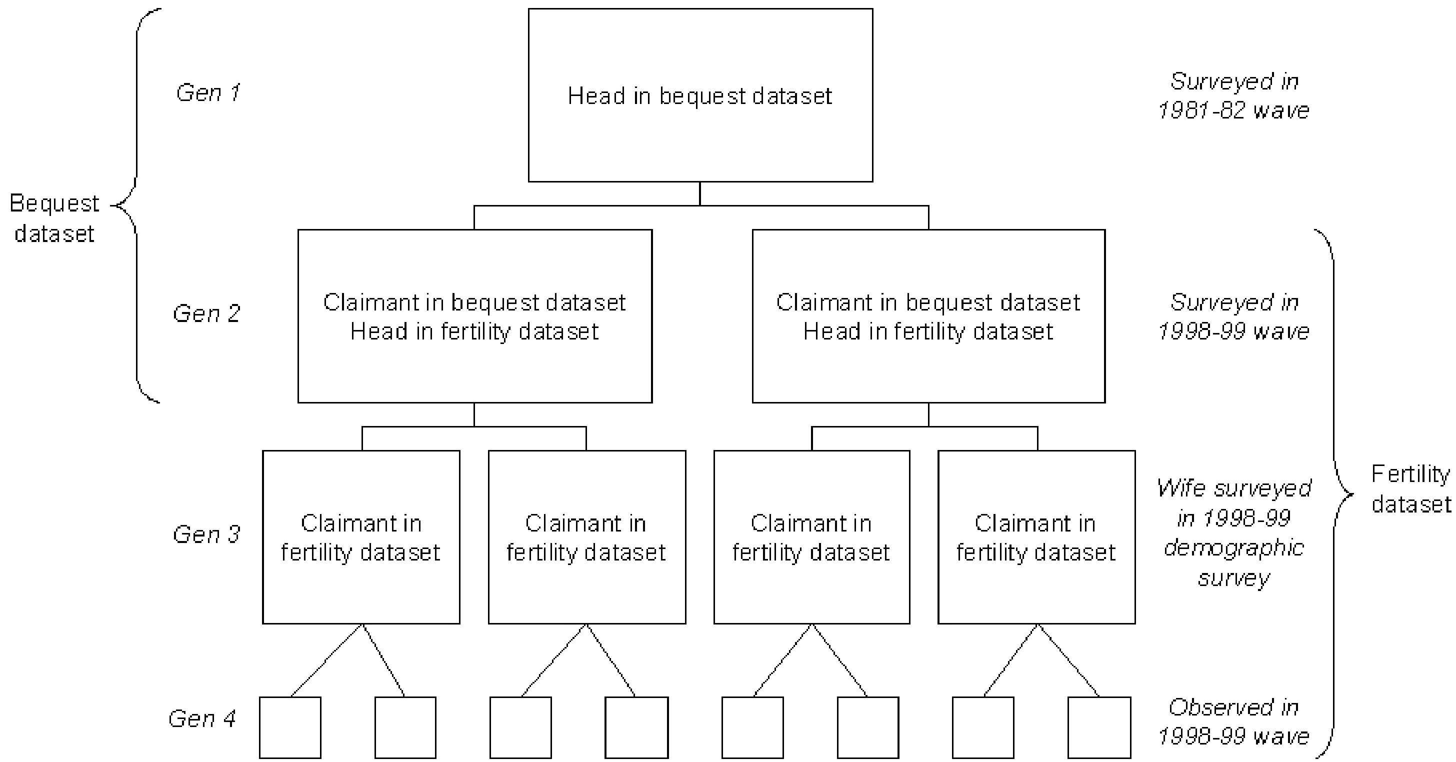

Figure 3: **Three generation family**

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)		**(II)**		**(III)**	
	Coeff.	**Std. Err.**	**Coeff.**	**Std. Err.**	**Coeff.**	**Std. Err.**
Number of own sons	**0.010** **	(0.005)	**0.023** ***	(0.008)	**0.018** ***	(0.007)
Number of own daughters	0.001	(0.005)	0.006	(0.008)	0.010	(0.007)
Other claimants' sons	**-0.012** ***	(0.003)	**-0.021** ***	(0.007)	**-0.018** ***	(0.004)
Other claimants' daughters	-0.004	(0.003)	**-0.017** ***	(0.007)	**-0.009** **	(0.004)
Split	-0.020	(0.019)	0.000	(0.028)	-0.016	(0.028)
Number of own sons * Split	0.003	(0.009)	-0.003	(0.015)	0.001	(0.013)
Number of own daughters * Split	0.014	(0.009)	0.008	(0.014)	0.026 *	(0.015)
Other claimants' sons * Split	-0.006	(0.006)	-0.008	(0.013)	-0.005	(0.007)
Other claimants' daughters * Split	0.001	(0.006)	0.009	(0.014)	-0.001	(0.007)
Birth Order	**-0.080** ***	(0.006)				
Left censored observations	35		8		27	
Uncensored observations	1019		362		657	
Right censored observations	13		6		7	

Notes: Double censored random effects tobit model. Standard errors are clustered at joint family level.
*** indicates coefficients are significant at 1% level. ** indicates coefficients are significant at 5% level.
* indicates coefficients are significant at 10% level. Source: REDS 1998-99.

Table 5: **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. *** indicates coefficients are significant at 1% level. ** indicates coefficients are significant at 5% level. * indicates coefficients are significant at 10% level.
Source: REDS 1998-99.

Table 6: **Results of test of strategic fertility within joint families**

Dependent variable: Reported pregnancy		
	I: Probit marginal effects	**II: Probit random effects coefficients**
Number of own sons	**- 0.065** ***	**- 0.339** ***
	(0.007)	(0.026)
Number of own daughters	**- 0.022** ***	**- 0.130** ***
	(0.005)	(0.025)
Other claimants' sons	**0.008** ***	**0.040** ***
	(0.003)	(0.013)
Other claimants' daughters	- 0.005	- 0.028 *
	(0.003)	(0.016)
Number of own sons * Split	- 0.046 *	**-0.240** ***
	(0.026)	(0.091)
Number of own daughters * Split	- 0.020	- 0.098
	(0.017)	(0.071)
Other claimants' sons * Split	- 0.001	0.006
	(0.009)	(0.043)
Other claimants' daughters * Split	0.015	0.071
	(0.010)	(0.048)
Constant		- 0.124

Notes: Values in parentheses are standard errors. Standard errors are clustered at joint family level. *** indicates coefficients are significant at 1% level. * indicates coefficients are significant at 10% level. N = 7,522 in 599 joint families. Source: REDS 1998-99.

Table 7: **Marginal effects of within family test**

Dependent variable: Reported pregnancy							
	Other claimants' sons						
	0	1	2	3	4	5	6
Number of own sons	**- 0.060** *** (0.006)	**- 0.063** *** (0.006)	**- 0.066** *** (0.007)	**- 0.069** *** (0.007)	**-0.072** *** (0.008)	**-0.075** *** (0.009)	**-0.078** *** (0.010)
Number of own daughters	**-0.020** *** (0.005)	**-0.021** *** (0.005)	**-0.022** *** (0.005)	**-0.023** *** (0.006)	**-0.024** *** (0.006)	**-0.025** *** (0.006)	**-0.026** *** (0.007)
Other claimants' sons	**0.008** *** (0.002)	**0.008** *** (0.003)	**0.009** *** (0.003)	**0.009** *** (0.003)	**0.009** *** (0.003)	**0.010** *** (0.004)	**0.010** *** (0.004)
Other claimants' daughters	- 0.004 (0.003)	- 0.004 (0.003)	- 0.005 (0.004)	- 0.005 (0.004)	- 0.005 (0.004)	- 0.005 (0.004)	- 0.006 (0.004)
Number of own sons * Split	- 0.042 * (0.022)	- 0.044 * (0.023)	- 0.046 * (0.024)	- 0.049 * (0.025)	- 0.051 * (0.026)	- 0.053 * (0.027)	- 0.055 * (0.028)
Number of own daughters * Split	-0.019 (0.015)	-0.020 (0.016)	-0.021 (0.017)	-0.021 (0.018)	-0.022 (0.018)	-0.023 (0.019)	-0.024 (0.020)
Other claimants' sons * Split	-0.001 (0.008)	-0.001 (0.008)	-0.001 (0.009)	-0.001 (0.009)	-0.001 (0.009)	-0.001 (0.010)	-0.001 (0.010)
Other claimants' daughters * Split	0.014 (0.009)	0.015 (0.010)	0.016 (0.010)	0.016 (0.011)	0.017 (0.011)	0.018 (0.012)	0.019 (0.013)

Notes: Values in parentheses are standard errors and are clustered at joint family level. *** and * indicate coefficients are significant at 1% and 10% level respectively. Coefficients for marginal effects evaluated at the means of the independent variables. Marginal effects for Other claimants' sons > 6 not shown. N = 7,522 in 599 joint families. Source: 1998-99.

Table 8: **Strategic fertility in land owning and landless families**

Dependent variable: Reported pregnancy				
	I: Probit marginal effects		II: Probit random effects coefficients	
	A: Landless head	B: Landowning head	A: Landless head	B: Landowning head
Number of sons	**-0.092** ***	**-0.064** ***	**-0.481** ***	**-0.333** ***
	(0.035)	(0.007)	(0.164)	(0.027)
Number of daughters	-0.021	**-0.022** ***	-0.145	**-0.128** ***
	(0.028)	(0.005)	(0.153)	(0.025)
Other claimants' sons	0.023	**0.008** ***	0.110	**0.039** ***
	(0.021)	(0.003)	(0.113)	(0.013)
Other claimants' daughters	-0.003	-0.004	-0.051	-0.026
	(0.022)	(0.004)	(0.197)	(0.016)
Number of sons * Split	0.027	**-0.057** **	0.104	**-0.291** ***
	(0.053)	(0.026)	(0.278)	(0.101)
Number of daughters * Split	-0.041	-0.022	-0.236	-0.103
	(0.062)	(0.019)	(0.292)	(0.076)
Other claimants' sons * Split	0.006	-0.001	0.045	0.004
	(0.024)	(0.009)	(0.184)	(0.047)
Other claimants' daughters * Split	-0.015	0.017	-0.035	0.082
	(0.042)	(0.011)	(0.271)	(0.052)
Constant			-0.145	

Notes: Values in parentheses are standard errors. Standard errors are clustered at joint family level. *** indicates coefficients significant at 1% level. ** indicates coefficients significant at 5% level. N = 7,522 in 599 joint families. Source: REDS 1998-99.

Table 9: **Strategic fertility before and after head's death**

Dependent variable: Reported pregnancy		
	A: Before head's death	**B: 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)

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

Table 10: **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)
Constant	- 1.166	-0.214	

Notes: Values in parentheses are standard errors. Standard errors are clustered at joint family level. *** indicates coefficients are significant at 1% level. ** indicates coefficients are significant at 5% level. 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.