The Economic Journal, **110** (*July*), 738–758. © Royal Economic Society 2000. Published by Blackwell Publishers, 108 Cowley Road, Oxford OX4 1JF, UK and 350 Main Street, Malden, MA 02148, USA.

GENDER BIAS, CREDIT CONSTRAINTS AND TIME ALLOCATION IN RURAL INDIA*

Elaina Rose

This paper examines the impact of a child's gender on the time allocation of rural Indian households for the five-year period subsequent to its birth. A theoretical model generates predictions for the effect of the birth of a boy relative to a girl (i.e., the gender shock) on household behaviour when the household is liquidity constrained and when it is not. The results from the empirical analysis are consistent with the case in which poorer households are liquidity constrained and less poor households are not. The interpretation of the finding that women in both groups of households work less subsequent to the birth of a boy relative to a girl differs in these two cases.

Policymakers and researchers interested in issues relating to the well-being of children have expressed concern that, in certain parts of the developing world, girls receive significantly less resources than boys. For South Asia, where gender bias is believed to be particularly acute (Sen, 1990; Behrman, 1992, 1998), differential treatment of boys and girls within families can be attributed to the greater pecuniary returns that sons provide to their parents relative to daughters. Therefore, economic models designed to explain the inferior treatment of girls in India typically consist of a model of intrahousehold resource allocation which allows for differences in the financial returns from sons and daughters.[1]

For rural India, the differences in returns may arise from the dowry system, from differences in the returns to male and female labour, and from the typically patrilocal family structure. Because these returns are realised years, or perhaps decades after the period of investment, the problem of investing in children is inherently an intertemporal one, and parents' decisions will depend upon their ability to borrow against future returns in order to finance investments in their children's early years. Attention to issues such as the presence of, and imperfections in, credit markets is critical when examining families in rural areas of developing countries, where there is typically a dearth of formal financial institutions. While families and other informal institutions

* Previous versions of this paper were presented at the Northeast Universities Development Consortium Conference, the conference on Gender Bias in Developing Countries and seminars at UC-Berkeley, UC-Riverside, Harvard/MIT, IFPRI, RAND, VPI, the University of Washington and the World Bank. I am grateful to editor Timothy Besley, and to Will Dow, Andrew Foster, Leah Gutierrez, Shelly Lundberg, Anna Paulson, Bob Pollak, Gene Silberberg and two anonymous referees for helpful comments/discussions; to ICRISAT for providing the data; to the Center for Studies in Demography and Ecology at the University of Washington for funding, and to Juan Cruces for excellent research assistance.

[1] For example, Rosenzweig and Schultz (1983) emphasise the role of gender differentials in returns in their study of the determinants of gender differentials in mortality.

may, in part, fulfill the role of formal markets, concern remains that households, and particularly the poorest households, remain constrained.[2]

This paper presents a theoretical and empirical analysis of the effect of the birth of a boy relative to a girl on time allocation in rural Indian households in which the early years of a child's life are viewed as an investment period and the child yields returns to parents in later years. The theoretical analysis is based on an intertemporal model of intrahousehold resource allocation and allows for two cases: one in which the household faces binding constraints in the credit market, and one in which it does not. One result generated from the empirical analysis is that women in both subsamples (the landless and small farm, or the 'poorest' households, and the medium and large farm, or the 'less poor' households) work less after the birth of a boy relative to a girl. However, because the empirical results are consistent only with a model in which the poorest household are constrained in the credit market and the less poor households are unconstrained, the finding has a very different interpretation for the two groups.

Section 1 of this paper presents a brief review of studies of intrahousehold resource allocation which address intertemporal issues. Section 2 presents the model that generates predictions for the effect of the birth of a boy relative to a girl on time allocation by gender, as well as on household-level savings and consumption, for the constrained and the unconstrained cases. When the household is unconstrained, the predicted effects can be decomposed into income effects, which result from the increase in household wealth generated by the birth of a son rather than a daughter, and substitution effects, which reflect the response of time allocated to child care to the higher returns from investing in a son relative to a daughter. When the household is constrained, the predictions differ because the child's gender affects the (endogenous) shadow interest rate, which in turn generates further income and substitution effects.

The empirical specification is presented in Section 3. The approach involves examining the effect of the birth of a boy relative to a girl (i.e., the 'gender shock') on days worked by men and women in rural Indian households for the five years of the child's life for the two subsamples. Section 3 also discusses econometric issues that arise when estimating the impact of the gender shock on household behaviour in an environment in which gender bias exists. The panel data set used in the analysis is described in Section 4.

The results are presented and discussed in Section 5. For poorer households the gender shock results in a decline in male leisure, the opposite result is found for the less poor households. For both subsamples, women work less following the birth of a son rather than a daughter. The results for the landless and small farmers are consistent with the case in which they are constrained in

[2] The literature on savings and credit, and the role of liquidity constraints in rural areas of developing countries is extensive and rapidly expanding. See, for example, Gersovitz (1988), Rosenzweig and Wolpin (1993), Besley (1995) and Morduch (1995).

the credit market, while the results for the medium and large farm households are consistent with the case in which they are unconstrained. The results accord with other studies of rural Indian households that find that it is the poorest who are most likely to be liquidity constrained (Morduch, 1990; Binswanger and Rosenzweig, 1993; Rose, 1999*b*).

The results of this analysis highlight the importance of addressing intertemporal issues when evaluating the effect of a child's gender on household behaviour. Women work less after the birth of a son relative to a daughter, but the finding has a very different interpretation for the two groups. For the medium and large farm households, the accompanying fall in male days worked indicates that these households are not constrained in the credit market, and the fall in female labour supply is consistent with an income effect on leisure due to the increase in wealth generated by the birth of a son rather than a daughter. However, for the landless and small farmers, the reduction of days worked in response to the gender shock is accompanied by a nearly simultaneous fall in male leisure. This is consistent with the presence of binding constraints in the credit market. The apparent increase in female home time is not solely an income effect on leisure, it represents an investment of women's time into care of the child; male leisure falls in order to compensate for the temporary shortfall in income due to reduced female labour supply.

1. Literature on Intertemporal Models of the Household

The role of credit constraints in parents' investments in children has been recognised. The work of Behrman *et al.* (1995) emphasises that the presence of liquidity constraints will alter conclusions regarding the efficiency of investments in child schooling. Two empirical analyses of allocation of resources to children in developing countries allow for imperfect financial markets. Foster (1995) shows that child growth patterns for children in landless households in rural Bangladesh are influenced by credit market imperfections. Jacoby and Skoufias's (1997) finding that, in rural India, children's schooling is related to seasonality in income is consistent with the presence of incomplete financial markets. However, the role of a child's gender not a focus of either of these studies.

A few papers address intertemporal issues in studying gender bias in rural India. Browning and Subramaniam (1995) find that the birth of a boy relative to a girl is associated with an increase in household consumption in the subsequent year and interpret this finding in terms of a model with perfect credit markets. Deolalikar and Rose (1998) finds that parents' savings behaviour for the five years subsequent to a child's birth is affected by the sex of the child. Behrman (1988) finds that gender bias in allocation of nutrients to children is greater in the lean relative to peak seasons. Rose (1999) shows that female relative to male child mortality is higher for children born in years in which the household experiences adverse weather shocks. These last two findings are consistent with differential consumption smoothing behaviour of parents with respect to sons and daughters.[3]

The literature on gender bias (and the literature on intrahousehold allocation in general) has typically not considered how conclusions would be affected by alternative assumptions regarding imperfections in credit (and other) markets. This paper presents an example in which this distinction between whether or not a household has access to the credit market is crucial in interpreting gender effects.

2. Theory

Consider a household that consumes in each of two periods, works only in the first period, and experiences a birth in the beginning of the first period. Household income in the first period consists of exogenous income, Y_1, and earnings from male and female labour; in the second period household income consists of exogenous income, Y_2. Because the objective of this model is to capture the role of children, and particularly sons, as old-age security, I consider the case in which Y_2 is low, so households will want to transfer resources from the present to the future. Households can accomplish this in two ways: by saving (S) at interest rate (r) and by investing female time (R_F) in caring for the child in the first period, yielding return $\beta g(R_F)$ in the second period. The function $g(.)$ can be thought of as a production function for child quality, while β represents the financial return that parents receive from the child's quality in the second period. The household's utility function is:

$$\max U(C_1, C_2, L_M, L_F) \tag{1}$$

where C_1 is first period consumption, C_2 is second period consumption, L_M is male leisure and L_F is female leisure. $\beta g(R_F)$ is assumed to be increasing and concave, and the utility function is assumed to be additively separable and increasing and concave in all arguments.

When the household is unconstrained, its problem is to maximise (1) subject to the intertemporal budget constraint:

$$C_1 + \frac{C_2}{1+r} + W_M L_M + W_F L_F = Y_1 + \frac{Y_2}{1+r} + W_M \overline{L}_M + W_F \overline{L}_F + \frac{\beta g(R_F)}{1+r} - W_F R_F \tag{2}$$

where W_M and W_F are male and female wage rates, and $\overline{L}_M$ and $\overline{L}_F$ are the endowments of male and female labour. When the household is constrained

[3] Additionally, Parish and Willis (1993) find that Taiwanese children with older sisters tend to get more education than those with older brothers, consistent with the presence of credit constraints combined with an earlier age that daughters provide returns to parents relative to sons. Garg and Morduch (1998*a*, *b*) show that the gender composition of a sibship affects the resources allocated to children using data from Ghana. This is interpreted in a model in which households may be resource constrained - and these 'resource constraints' may be interpreted as credit constraints.

in the credit market, it maximises (1) subject to (2) and the additional constraint:

$$S = Y_1 + W_M(\overline{L}_M - L_M) + W_F(\overline{L}_F - L_F - R_F) \geqslant 0. \qquad (3)$$

The effect of the gender shock on C_1, L_M, L_F, R_F and S can be captured by an increase in β on each of these endogenous variables. When the household is unconstrained, the total effect can be decomposed into income and substitution effects, where the income effect arises from the increase in wealth which is realised in the second period, and the substitution effect arises because a higher level of β increases the marginal return to transferring resources into the future by investing female (most likely maternal) time in the child.

When the household is constrained there are additional effects that arise through the endogenous shadow interest rate. While the effects of changes in β on consumption leisure, time in child care, and savings are ambiguous *a priori* individually, it is possible to look at the set of total effects as a whole in order to differentiate the constrained case from the unconstrained case.

Comparison of the set of predictions with the results from the empirical analysis also provides a basis for determining whether the increase of female home time to in response to the gender shock represents an increase in leisure (i.e., L_F) or an increase in time allocated to the child (i.e., R_F). The signs of the income effects, substitution effects, and the effects associated with a change in the shadow interest rate resulting from an increase in β in the constrained and unconstrained cases are summarised in Table 1.

Case I: Household Unconstrained in Credit Market

When the household is unconstrained, or when the constraint $S \geqslant 0$ is not binding, the first order conditions imply that $U_3/U_4 = W_M W_F$, and that

Table 1

Predicted Effects of an Increase in β on Endogenous Variables

Total Effects for:
Unconstrained Case (Column 3)
Constrained Case A (Column 5)
Constrained Case B (Column 6)

	(1) Income effect	(2) Subst. effect	(3) Total effect (Unconst.) (Case I)	(4) Shadow interest rate effect	(5) Total effect (Const.) (Case IIA)	(6) Total effect (Const.) (Case IIB)
C_1	+	0	+	−	+	−
L_M	+	0	+	−	+	−
L_F	+	0	+	−	+	−
R_F	0	+	+	−	−	+
$L_F + R_F$	+	+	+	−	−	+
S	−	−	−	+	0	0

$U_1/U_2 = \beta g'(R_F)/W_F = 1 + r$. Because leisure and consumption are normal under the maintained assumptions, an increase in β results in positive income effects for C_1, L_M and L_F, although there are no substitution effects because the presence of the credit market renders the investment and consumption decisions separable. A positive substitution effect of an increase in β on R_F arises because increasing β increases the returns to investing in children relative to saving at the market interest rate. Both the income and substitution effects lead to a reduction in savings following the birth of a boy relative to a girl. In this case, the total effect of an increase in β on each endogenous variable is the sum of the income and substitution effects. The signs of these income, substitution, and total effects are reported in columns (1), (2) and (3) of Table 1, respectively.

Case II: Household Constrained in the Credit Market[4]

The Kuhn-Tucker conditions when the constraint $S \geqslant 0$ is binding imply that $U_3/U_4 = W_M/W_F$, and that $U_1/U_2 = \beta g'(R_F)/W_F > 1 + r$.[5] In this case, in addition to the possible income and substitution effects of a change in β on the set of endogenous variables, there are additional effects which arise through the endogenous shadow interest rate, r^*. The shadow interest rate is the market interest rate at which the household, if unconstrained, would choose to set savings exactly to zero. When the household is constrained, the investment and consumption decisions are no longer separable.

The effects of an increase in β on consumption and male and female leisure each consist of an income effect, which corresponds to the income effect in the unconstrained case, and the shadow interest rate effects. I will illustrate this in terms of the effect of an increase in β on C_1. The expressions and intuition for the effects on male and female leisure are analogous.

When the household is constrained, the (Marshallian) demand for consumption in the first period can be expressed as:

$$C_1 = C_1[W_M, W_F, r^*(.), Y(.)] \tag{4}$$

where $r^*(.) = r^*(W_M, W_F, Y_1, Y_2, \beta)$ and $Y(.)$ is lifetime full income in terms of period 1 consumption, i.e., the right hand side of (2), evaluated at $r = r^*$.

Differentiating (4) with respect to β yields:

[4] This problem is similar to Strauss's (1986) analysis of the effects of an increase in an output price on the behaviour of an agricultural household when the labour market is absent.

[5] In the case in which there is a binding upper bound on borrowing, i.e., a negative binding lower bound on savings, the model could be respecified by allowing for $S > \overline{S}$, with $\overline{S} < 0$ in (3). This does not change any of the signs of the predicted effects, although footnote 7 describes a difference in interpretation of the effects.

$$\frac{\partial C_1}{\partial \beta} = \left.\frac{\partial C_1}{\partial \beta}\right|_{r^*} + \frac{\partial C_1}{\partial r^*}\frac{\partial r^*}{\partial \beta} \tag{5}$$

$$= \left.\frac{\partial C_1}{\partial Y}\right|_{r^*}\frac{\partial Y}{\partial \beta} + \left[\left(\left.\frac{\partial C_1}{\partial r^*}\right|_{Y} + \left.\frac{\partial C_1}{\partial Y}\right|_{r^*}\frac{\partial Y}{r^*}\right)\frac{\partial r^*}{\partial \beta}\right]. \tag{6}$$

The first term in (6) is a positive income effect corresponding to the income effect in the unconstrained case. The bracketed term contains a set of terms which constitute the decomposed shadow interest rate effect.[6]

In this model it is possible to sign the shadow interest rate effect by examining its components. Because an increase in r^* represents an increase in the price of period 1 consumption,

$$\left.\frac{\partial C_1}{\partial r^*}\right|_{Y} < 0.^7$$

Because consumption is normal,

$$\left.\frac{\partial C_1}{\partial Y}\right|_{r^*} > 0.$$

When the shadow interest rate increases, the value of the endowment denominated in terms of period 1 consumption falls, so

$$\frac{\partial Y}{\partial r^*} < 0.$$

Finally, an increase in β increases both lifetime full income and the returns to investment of women's time in care of the child. The first effect leads to an increase in the demand for consumption and leisure in the first period because they are normal, and the second effect leads to an increase in R_F. Because these increases in demands must all be financed in period 1, pressure on the household to borrow increases when β increases, so

$$\frac{\partial r^*}{\partial \beta} > 0.$$

The effect of an increase in β on R_F consists of a substitution effect corresponding to the substitution effect in the unconstrained case and a shadow interest rate effect. The demand for R_F can be expressed as:

$$R_F = R_F[W_F, \beta, r^*(.)]. \tag{7}$$

[6] The shadow interest rate effects could also be decomposed further into income and substitution effects of the shadow interest rate on consumption, and the shadow interest rate itself can be decomposed into 'income and substitution effect-like' terms. See Strauss (1986).

[7] Since the household does not borrow or lend when constrained in this model there is no income effect on C_1 resulting from a change in r^*, only a negative substitution (compensated price) effect. However, if the model were modified to allow for $S > \overline{S}$, with $\overline{S} < 0$, then

$$\left.\frac{\partial C_1}{\partial r^*}\right|_{Y}$$

would also contain a negative income effect which would reinforce the substitution effect.

The change in R_F resulting from a change in β is:

$$\frac{\partial R_F}{\partial \beta} = \left.\frac{\partial R_F}{\partial \beta}\right|_{r*} + \frac{\partial R_F}{r^*}\frac{\partial r^*}{\partial \beta}. \tag{8}$$

The first term is the substitution effect, which is positive. The second term is the negative shadow interest rate effect.

In summary, the effects of an increase in β on C_1, L_M, L_F, and R_F can be decomposed into positive income and substitution effects, corresponding to the income and substitution effects in the unconstrained case and shadow interest rate effects, which are negative. These signs of the income, substitution and shadow interest rate effects are reported in columns (1), (2), and (4) of Table 1, respectively.

Although the predicted effects of an increase in β on C_1, L_M, L_F, and R_F are ambiguous *a priori*, it is shown in Appendix 1 that the effect of an increase in β on total home time (i.e., the sum of L_F and R_F) is dominated by the effect on R_F. Additionally, the constrained case can be divided into two sub-cases. In Case II.A (Case II.B) the income effects on consumption and the leisures outweigh (are outweighed by) the shadow interest rate effects on these variables, and the shadow interest rate effect on R_F outweighs (is outweighed by) the substitution effect. Also, the total effect of an increase in β on total home time is dominated by the effect on R_F. The signs of the total effects in Case II.A and Case II.B are listed in columns (5) and (6) of Table 1, respectively.

The set of predictions generated from the models can be used to distinguish households that are constrained from households that are unconstrained using data on time allocation for males and females. When the household is unconstrained, both male leisure and female home time increase in response to the birth of a boy relative to a girl. When the household is constrained, either male leisure will increase and female home time will fall (Case II.A.) or male leisure will fall and female home time will increase (Case II.B.).

Note that female home time could increase after the birth of a boy relative to a girl if the household is unconstrained (Case I) or if the household is constrained (Case II.B). If the household is unconstrained this represents an increase in both leisure and time allocated to the child, if the household is constrained, in Case II.B, then female leisure is *lower* and the increase in home time represents only an increase in time allocated to care of the child.

3. Empirical Specification and Econometric Issues

The effect of the birth of a son relative to a daughter within the first five years of the child's birth can be captured using the following empirical specification:

$$DAYSWORK_{kit} = \alpha_{ki} + v_{kt} + \beta_{k,04}\sum_{l=0}^{4} Boy_{i,t-l} + \gamma_{k,04}\sum_{l=0}^{4} Girl_{i,t-l} + \boldsymbol{\theta}\mathbf{X}_{it} + \epsilon_{it} \tag{9a}$$

where $DAYSWORK_{kit}$ is total days worked by individuals in category k ($k \epsilon \{M, F\}$) in household i in period t, α_{ki} is a household fixed effect, v_{it} is a village × time fixed effect, $Boy_{i,t-l}$ ($Girl_{i,t-l}$) is the number of boys (girls) born to household i in year $t-l$, $\sum_{l=0}^{4} Boy_{i,t-l}$ ($\sum_{l=0}^{4} Girl_{i,t-l}$) is the *total* number of boys (girls) born within the past five years, and $\beta_{k,04}$ ($\gamma_{k,04}$) is the effect of the birth of a boy (girl) born at any time within the past five years on days worked for category k. A set of controls for household wealth and family size are included in the vector X. The effect of the birth of a boy relative to a girl based on estimates of (9a) is the difference, $\delta_{k,04} = \beta_{k,04} - \gamma_{k,04}$. I will refer to this difference as 'the impact of the gender shock'. A test for the significance of the gender shock in this specification is a test of the hypothesis $\delta_{k,04} = 0$.

It may be the case the effect of the birth of a child is different in the year of its birth relative to later years. These effects can be distinguished using the following specification:

$$DAYSWORK_{kit} = \alpha_{ki} + v_{kt} + \beta_{k,0} Boy_{it} + \gamma_{k,0} Girl_{it} + \beta_{k,14} \sum_{l=1}^{4} Boy_{i,t-l} + \gamma_{k,14} \sum_{l=1}^{4} Girl_{i,t-l} + \theta X_{it} + \varepsilon_{it}. \qquad (9b)$$

In a more general specification, the effect of the birth of a boy and girl in each of the five years can be entered separately:[8]

$$DAYSWORK_{kit} = \alpha_{ki} + v_{kt} + \sum_{l=0}^{4} \beta_{kl} Boy_{i,t-l} + \sum_{l=0}^{4} \gamma_{k,l} Girl_{i,t-l} + \theta X_{it} + \varepsilon_{it}. \qquad (9c)$$

In (9c) β_{kl} (γ_{kl}) are parameters representing the effect of the birth of a boy (girl) l years ago and the effect of the birth of a child l years ago is captured by the difference $\delta_{kl} = \beta_{kl} - \gamma_{kl}$.

Each specification also includes a set of controls for household size and wealth. Additionally, all fixed household characteristics are subsumed in the household fixed effect, all village level time-varying variables, such as prices and aggregate shocks are subsumed in the village × time fixed effect.

Estimates of the β's and the γ's individually from (9a) to (9c) may be biased due to the endogeneity of fertility, however estimates of the δ's will be unbiased if births are reported correctly, if the sex of a child does not affect the parents' subsequent fertility decision, and if the standard assumptions underlying this fixed effects model are otherwise satisfied.

There are two potential sources of bias that need to be considered in estimating gender effects using this approach in an environment in which gender bias exists. The first arises when preference for boys is manifest through parents' having a minimum or target number of sons and fertility is also correlated with unobservables in the estimating equation, such as parents' preferences or potential market productivity. The implications are illustrated

[8] Deolalikar and Rose (1998) use a similar empirical specification to estimate the impact of the gender shock on savings, income and consumption.

in Appendix 2. The estimated gender shock will be unbiased for earlier children (i.e., children born prior to the point at which subsequent fertility depends upon the sex composition of the existing children), but biased for later children.

The second potential source of bias has to do with the treatment of children who are born and die within the sample period. Because gender differentials in mortality are, in some part, endogenous, the data are not adjusted when a child is reported to have died by age four. That is, gender differences in mortality (or gender differences in allocation of resources that result in gender differences in mortality) are one means that households employ *ex post* in order to mitigate the effect of an adverse gender shock.

A third potential source of bias arises when the sex of the child is correlated with the error term in the estimating equation because the propensity to report births of boys differs from the propensity to report girls. This could arise under two circumstances. First, if the mortality of girls is higher than the mortality of boys in the first few weeks of life and births of children who die in the first few weeks are not reported on the household roster. Second, this could arise without actual sex-differentials in mortality if the birth of a boy is considered a more significant event for a family than the birth of a girl, and girls who die shortly after birth are less likely to be reported to have ever been born than boys. Both of these circumstances may arise in an environment in which gender bias exists.[9]

It is not feasible to employ an instrumental variables technique to correct for bias due to differential reporting of births of boys and girls; however, it is possible to sign the bias. The household fixed effects in (9*a*) to (9*c*) control for correlation of regressors with unobservable permanent household characteristics, such as parents' education or preferences, and the village $\times$ time fixed effects control for correlation of regressors with unobservable transitory aggregate (village level) characteristics such as bad weather. Therefore the estimates could only be biased if the child's sex is correlated with idiosyncratic transitory factors. Labour supply in rural India has been shown to increase in response to adverse shocks (Rose, 1999*a*; Kochar, 1999), and the survival probabilities of girls relative to boys has been shown to be higher in years in which there were favourable weather shocks (Rose, 1999*b*). Therefore, the sign of the potential bias from this source on estimates of δ would likely be positive.

[9] Note that in these circumstances, in any individual-level estimating equation in which a child's sex is an independent variable, its coefficient is subject to bias due to potential correlation of the variable with unobservables in the estimating equations. (Rose, 1999b). Similarly, the difference in the coefficients on the number of boys and the number of girls in a household in an equation explaining some household level outcome might also be biased.

4. Data

This analysis requires data on time allocation for males and females in the household and on births and deaths of children in the ages corresponding to period 1 in the model. The data used in this analysis were obtained from the International Crop Research Institute for the Semi-Arid Tropics (ICRISAT) of India. Two types of data are required. First, time allocation data are required. This information was obtained from schedule 'K', which records number of days worked in own-farm production, farm employment, non-farm employment, and government employment, by individual. The files do not record time spent in non-farm self-employment. Individual level data on days worked were aggregated to the household level, and days worked by non-family members are excluded from the household total. Consistent data were available for three villages (Aurepalle, Shirapur and Kanzara) and six crop years (1979–84).

Second, data on household demographics are required. This information is obtained from schedule 'C' which is designed to record all household members by gender and age in 1979 as well as births of children by sex for the ages corresponding to period one of the model. This period was chosen to be the first five years of the child's life for two reasons. First, this is roughly the period in which the mother's time input is most important for the child and in which the child is not yet productive. Second, this is the longest period for which it is possible to collect data on births and deaths corresponding to the years in which the time allocation data are available.

The original ICRISAT sample was stratified into four groups: landless households, and small, medium and large farm households. Following Morduch (1990) and others, the analysis is performed on two groups: the landless and small farmers (poorer households) and the medium and large farmers (less poor households). Measures of household net worth and the number of males and females in the household, obtained from the ICRISAT summary files, are included as controls. Sample statistics are reported in Appendix 3.

Examination of the data suggests that the problem of under-reporting of births, as discussed at the end of the previous section, is an issue in this data set: There are 88 boys and 70 girls reported to have been born between 1975 and 1984 which implies a reported sex ratio at birth[10] of 126 rather than the expected ratio of approximately 105. Only one child is reported to have died in the first year, which is the year in which child mortality would be expected to be highest.[11] However, the under-reporting is concentrated in the medium and large farm households, where 50 boys and 35 girls are reported to have been born; the corresponding numbers for the landless and small farmers are 35 girls and 38 boys.

[10] The sex ratio at birth is the number of births of boys divided by the number of births of girls, times 100.

[11] Eleven deaths are reported for children ages 1 to 4.

5. Results

The results for women (men) in medium/large farm households are reported in Table 2 (3); the results for women (men) in landless and small farm households are reported in Table 4 (5). The structures of Tables 2 to Table 5 are identical. There are three panels in each table in which the results from (9*a*) to (9*c*) are reported. The left portion of each table presents the estimated parameters and the t-statistics associated with the tests that the respective parameters are zero. Additional diagnostic statistics are reported on the right hand side of each table. For the second two specifications, the p-values associated with the test that the effects of the birth of a boy relative to a girl for each lag are jointly significant (i.e., $p(\delta_L = 0, \forall L)$) and the test that the effects for all lags are the same are reported. For (9*b*), the test is that $\delta_0 = \delta_{14}$. For (9*c*), the test in the row corresponding to lag L is the test that $\delta_0 = \delta_1 = \delta_2 = \delta_3 = \delta_4$. For (9*c*) the p-value associated with the test that the coefficients for the adjacent lags are reported $\delta_L = \delta_{L-1}$ in the row corresponding to lag L.

Table 2 presents the results for women in medium and large farm households. When the five years are pooled together in (9*a*), the estimate of the gender shock is negative and significant. When the first year is broken out from the grouping in (9*b*), the estimated gender shock is insignificant for the year of birth, but remains significant for the subsequent years grouped

Table 2

*Estimated Effect of Birth of Boy, Girl, and Difference Females, Medium and Large Farm Households**

(N= 331, t-statistics in parentheses)

Equation	Lag	Diff. (δ_L)	Boy (β_L)	Girl (γ_L)	R^2	p ($\delta_L = 0, \forall L$)	p ($\delta_L = \delta_M, \forall L \neq M$)	p ($\delta_L = \delta_{L-1}$)
(9*a*)	0–4	−29.5	1.1	30.1	0.74			
		(1.9)	(0.09)	(2.4)				
(9*b*)	0	−26.9	4.2	31.1	0.74	0.19	0.93	
		(0.9)	(0.15)	(1.8)				
	1–4	−29.8	0.61	30.4				
		(1.8)	(0.05)	(2.2)				
(9*c*)	0	−10.5	8.6	19.1	0.74	0.25	0.44	
		(0.31)	(0.32)	(1.0)				
	1	−16.5	−3.4	13.1				0.87
		(0.7)	(0.2)	(0.7)				
	2	−1.8	13.1	15.0				0.62
		(0.07)	(0.9)	(0.7)				
	3	−38.4	18.4	56.8				0.20
		(1.5)	(1.0)	(2.9)				
	4	−47.7	−21.0	26.7				0.44
		(2.1)	(1.6)	(1.3)				

* Household fixed effects estimates, with Huber standard errors at household level. Controls for total number of males, total number of females, household net worth, and village × time dummy variables are included.

together. Breaking the grouping down further by estimating effects for each lag independently in (9*c*), indicates that the effect is significant only for lag 4 individually. However, because the standard errors are large, it is impossible to reject hypotheses that the effects for adjacent lags are equal ($p = 0.87$, 0.62, 0.20, 0.44) or that all five effects are equal ($p = 0.44$). In this case, it is possible to conclude that the birth of a boy relative to a girl reduces labour supply for these households and that the effect is significant for $L = 4$, but it is not possible to say more regarding the timing of the effects.

The corresponding results for men in medium and large households are reported in Table 3. Here, again, there is a significant negative effect when lags 0 to 4 are pooled together in (9*a*) and when lags 1 to 4 are pooled together in (9*b*). In (9*c*), when the effects are estimated for each of the lags separately, the estimated effects are all negative and, with the exception of $L = 1$, they are all statistically significant. In three of the four cases, the hypothesis that the effects for adjacent lags are the same cannot be rejected (the exception being $L = 1$ vs. $L = 2$). Comparing the results with the predictions from the theoretical model which are summarised in Table 1 indicates that these results for the medium and large farm households are consistent only with the case in which the households are unconstrained in the credit market.

Table 4 reports the results for women in the landless and small farm households. The estimated effect of the birth of the boy relative to a girl is insignificant for the pooled specification in (9*a*). The effect for $L = 0$ is significant ($t = 1.7$) in (9*b*) and it is possible to reject the hypothesis that

Table 3

*Estimated Effect of Birth of Boy, Girl, and Difference Males, Medium and Large Farm Households**

(N = 331, t-statistics in parentheses)

Equation	Lag	Diff. (δ_L)	Boy (β_L)	Girl (γ_L)	R^2	p ($\delta_L = 0$, $\forall L$)	p ($\delta_L = \delta_M$, $\forall L \neq M$)	p ($\delta_L = \delta_{L-1}$)
(9*a*)	0–4	−64.0	−18.8	45.2	0.82			
		(2.9)	(1.4)	(2.5)				
(9*b*)	0	−51.9	−6.8	45.1	0.82	0.01	0.64	
		(1.5)	(0.3)	(1.5)				
	1–4	−65.4	−20.7	44.7				
		(3.0)	(1.5)	(2.4)				
(9*c*)	0	−69.5	−9.3	60.2	0.82	0.01	0.38	
		(1.9)	(0.4)	(2.1)				
	1	40.4	9.4	49.8				0.47
		(1.0)	(0.5)	(1.6)				
	2	−110.1	−44.3	65.8				0.07
		(3.4)	(2.9)	(2.4)				
	3	−56.3	−37.4	18.8				0.13
		(2.1)	(2.1)	(0.7)				
	4	−64.7	−8.3	56.5				0.82
		(2.2)	(0.4)	(2.1)				

* Notes as Table 2.

Table 4

*Estimated Effect of Birth of Boy, Girl, and Difference Females, Landless and Small Farmers**

(N= 344, t-statistics in parentheses)

Equation	Lag	Diff. (δ_L)	Boy (β_L)	Girl (γ_L)	R^2	p ($\delta_L = 0$, $\forall L$)	p ($\delta_L = \delta_M$, $\forall L \neq M$)	p ($\delta_L = \delta_{L-1}$)
(9*a*)	0–4	−0.55	−25.0	−24.5	0.63			
		(0.03)	(1.7)	(1.2)				
(9*b*)	0	−56.7	−54.7	2.0	0.63	0.11	0.04	
		(1.7)	(2.2)	(0.07)				
	1–4	10.0	−19.2	−29.1				
		(0.5)	(1.3)	(1.3)				
(9*c*)	0	−70.6	−41.6	29.1	0.65	0.16	0.11	
		(1.9)	(1.7)	(0.9)				
	1	12.0	16.8	4.7				0.07
		(0.3)	(0.6)	(0.15)				
	2	17.8	4.3	−13.4				0.89
		(0.5)	(0.15)	(0.36)				
	3	−14.0	−6.8	7.0				0.45
		(0.4)	(0.33)	(0.2)				
	4	45.7	−28.4	−74.2				0.14
		(1.5)	(1.3)	(3.0)				

* Notes as Table 2.

$\delta_0 = \delta_{14}$ for (9*b*) (p = 0.04). In (9*c*) when the lags are entered separately, the effect for $L = 0$ is significant (t = 1.9) although the estimates are not significant for other lags. The hypothesis that $\delta_0 = \delta_1$ can be rejected (p = 0.07), although it is not possible to reject the hypothesis that the effects for other pairs of adjacent lags are equal. In summary, women in the poorer households work less in the year of the child's birth, and the effect of the gender shock for the year of birth is significantly less than the effect for the other years pooled.

The corresponding results for men are reported in Table 5. The results for (9*a*) indicate that men work more in response to the birth of a boy vs. a girl. The hypotheses that $\delta_0 = \delta_{14}$ and is rejected. The results from (9*c*) indicate that the hypothesis that the effects for all lags, and the hypotheses that $\delta_0 = \delta_1$ and $\delta_1 = \delta_2$, (p = 0.02, 0.001, 0.08, respectively) can be rejected, although it is not possible to reject the hypotheses that the effects are the same for other pairs of adjacent years. The results when the estimate of the gender shock is estimated for each of the lags individually in (9*c*) indicate that this increase occurs only in the year subsequent to the year of the child's birth.

The finding that women in the landless and small farm household work less after a boy is born, and that the men work more is consistent with the Case II.B from Table 1 in which households are constrained in the credit market. The timing of the effects is consistent with a story about credit constraints: women work less in the year in which a boy is born relative to a girl, and men work more in the subsequent year in order to compensate for the shortfall in income.

These results can be combined with the estimated impacts of the gender of

Table 5

*Estimated Effect of Birth of Boy, Girl, and Difference Males, Landless and Small Farmers**

(N = 344, t-statistics in parentheses)

Equation	Lag	Diff. (δ_L)	Boy (β_L)	Girl (γ_L)	R^2	p ($\delta_L = 0$, $\forall L$)	p ($\delta_L = \delta_M$, $\forall L \neq M$)	p ($\delta_L = \delta_{L-1}$)
(9*a*)	0–4	48.7	11.3	−37.4	0.68			
		(2.1)	(0.7)	(2.2)				
(9*b*)	0	−8.5	−30.1	−21.6	0.69	0.01	0.03	
		(0.25)	(1.6)	(0.7)				
	1–4	58.4	19.8	−38.6				
		(2.5)	(1.2)	(2.1)				
(9*c*)	0	−28.4	−27.4	1.1	0.70	0.01	0.02	
		(0.8)	(1.3)	(0.03)				
	1	110	52	−58.3				0.001
		(3.3)	(.19)	(2.2)				
	2	42.2	84.7	42.5				0.08
		(1.2)	(2.8)	(1.5)				
	3	1.4	31.5	−23.2				0.80
		(1.3)	(1.2)	(0.7)				
	4	38.2	−14.3	−52.5				0.02
		(1.3)	(0.7)	(2.3)				

* Notes: as Table 2.

a child at birth on consumption and savings reported in Browning and Subramaniam (1995) (B&S) and Deolalikar and Rose (1998) (D&R).[12] The entire set of results are summarised in Table 6 and can be compared to the predicted effects of the gender shock reported in Table 1. The decrease in savings and increase in consumption for the medium and large farm households is consistent with the model in which households are unconstrained in the credit market. The absence of an effect of the birth of a boy relative to a girl on the gender shock which is found for the landless and small farm households is consistent with a model in which the households are constrained in the credit market.[13] Note that since labour supply of men and women in both types of households responds to the shock, the results are not consistent with a model in which households are constrained in the labour market.

5.1. *The Role of Preferences*

This analysis has focused on the role of differential returns as the mechanism through which a child's gender affects household behaviour. In this section I

[12] Deolalikar and Rose estimate the effect of the gender shock on savings for landless/small farm households and medium/large farm households, and on consumption and income for the medium/large farmers with the same empirical specification used here. Browning and Subramaniam follow a similar approach, which is essentially to estimate the effect of the shock on consumption for $L = 1$ only, in the case in which $\beta_1 = -\gamma_1$, for the two groups.

[13] It is not possible to reject the hypothesis that there is no effect of the gender of a child on consumption for the poorer households. This may be due to the difficulty in measuring consumption for rural households.

Table 6

*Summary of Results on Impact of Gender Shock (Boy vs. Girl)**

	Source	Landless and small farmers	Medium and large farmers
Days worked by females		<0 $(L=0)$	<0 $(L=4)$
Days worked by males		>0 $(L=1)$	<0 $(L=0, 2, 3, 4)$
Savings (S)	D&R	<0 $(L=0–4)$	0
Consumption (C_1)	D&R, B&S	>0 $(L=1)$	0

* D&R: Deolalikar and Rose (1998); B&S: Browning and Subramaniam (1995); otherwise, the source of the results is this paper.

consider the predicted effects of the gender shock that would arise under two versions of a pure preference model and compare these predictions with those obtained from the investment model presented in Section 2.

In a unitary household model, if parents prefer sons to daughters (i.e., if $\beta g(R_F)$ appears in the utility function, where β is higher for boys than for girls because parents care more about outcomes of sons than outcomes of daughters) then R_F will be higher following the birth of a son relative to a daughter. If there are no differences in returns as well, consumption of goods and leisure must fall in order to accommodate the increased resources devoted to the son rather than the daughter. When the household is constrained the reductions must appear simultaneously or shortly following the increase in R_F, when the household is unconstrained, then the reductions will be spread over a long time horizon and may not appear in the first five years of the child's life. The increased consumption of male leisure subsequent to the gender shock for the medium and large farmers is inconsistent with this model. The increase in women's home time for the landless and small farmers is consistent with this version of a pure preference model. However, as with the investment model, the nearly simultaneous increase in male labour supply implies that these poorer households are constrained and that the increase in female home time arises through investment of time into the child rather than an increase in leisure.

Second, in a bargaining model, it may be the case that the birth of a boy increases the mother's status within the household. In this case, her leisure may increase subsequent to the birth of a son relative to a daughter; if she cares more about child well-being than do other decision-makers within the household (Thomas, 1990), then R_F may increase as well. In both cases, if there is no differential in returns, then consumption and male leisure must fall in order to finance the reduction in female time worked. The result that male leisure increases subsequent to the gender shock which is found for the medium and large farmers is inconsistent with this model. The increase in female home time in response to the gender shock for the landless and small farmers is consistent with the bargaining model; again, the subsequent drop in male leisure suggests that the household is credit constrained.

In summary, the increase in male leisure which is found for the medium and large farmers rules out a pure preference explanation for the findings of the empirical analysis. For the landless and small farmers, the pattern is consistent with the two preference models – but both imply that these households are liquidity constrained. Also, in the unitary model, and in the bargaining model in which mothers care more about children than other decision-makers, the results imply that boys receive more of their mothers' time than do girls.[14]

6. Conclusion

This paper has addressed the intertemporal nature of the problem of gender bias in India in terms of a two-period model of intrahousehold resource allocation and estimated the effect of the birth of a son relative to a daughter on household time allocation. The sample of households from rural India used in the empirical analysis is divided into two groups: the landless and the small farmers, and the medium and large farmers. The results, in combination with estimates from previous work, are consistent with a model in which the former group faces binding constraints in the credit market, while the latter does not.

The findings in this paper underscore the importance of considering intertemporal issues when testing for gender bias in South Asia. Interpretation of the reduction in days worked by women in response to the gender shock differs substantially for the two subsamples due to the finding of binding credit constraints for one, but not for the other. For medium and large farm households, the increase in home time is consistent with an income effect: when a son is born, the family is richer and its women work less and consume more true leisure. In contrast, the decline in male leisure which accompanies the increase in apparent female leisure for the landless and small farmers suggests that the decline in female time worked in response to the gender shock is attributable to substitution of the mother's time from other productive activities into care of her child. In this case, increased apparent female leisure reflects time devoted to the care of a male infant which would not be allocated to the child if he were a girl.

University of Washington

Date of receipt of first submission: January 1997
Date of receipt of final typescript: October 1999

Appendix 1. Computing Predicted Effects from an Increase in β on Endogenous Variables

Case I: Household Unconstrained in the Credit Market

When the household is unconstrained its problem is

$$\max U(C_1, C_2, L_M, L_F) \tag{A1.1}$$

[14] Additionally, if parental preference for sons may be manifested through a minimum or target number of sons, fertility will be lower subsequent to the birth of a son rather than a daughter. Under certain conditions, this will result in biased coefficients; this is discussed in Section 3.

$$\text{s.t.}\, C_1 + \frac{C_2}{1+r} + W_M L_M + W_F L_F = Y_1 + \frac{Y_2}{1+r} + W_M \overline{L}_M + W_F \overline{L}_F + \frac{\beta g(R_F)}{1+r} - W_f R_F.$$

Assuming that $g(.)$ is increasing and concave ($g' > 0$, $g'' < 0$), the utility function is additively separable and increasing and concave in each argument ($U_k > 0$, $U_{kk} < 0\ \forall k$; $U_{kk} < 0\ \forall j \neq k$), and assuming an interior solution, the following comparative statics results are obtained from the above problem:

$$\frac{\partial C_1}{\partial \beta} = g\frac{\partial C_1}{\partial Y_2} = \frac{g\rho^2 U_1 U_{22} U_{33} U_{44} \beta g''}{-\Delta} > 0$$

$$\frac{\partial L_M}{\partial \beta} = g\frac{\partial L_M}{\partial Y_2} = \frac{g\rho^2 W_M U_1 U_{11} U_{22} U_{44} \beta g''}{-\Delta} > 0$$

$$\frac{\partial L_F}{\partial \beta} = g\frac{\partial L_F}{\partial Y_2} = \frac{g\rho^2 W_F U_1 U_{11} U_{22} U_{33} \beta g''}{-\Delta} > 0$$

$$\frac{\partial R_F}{\partial \beta} = \frac{\rho g' U_1 [U_{11} U_{22} (W_F^2 U_{33} + W_M^2 U_{44}) + \rho^2 U_{11} U_{33} U_{44} + U_{22} U_{33} U_{44}]}{-\Delta} > 0$$

$$\frac{\partial (L_F + R_F)}{\partial \beta} = \frac{\partial L_F}{\partial \beta} + \frac{\partial R_F}{\partial \beta} > 0$$

$$\frac{\partial S}{\partial \beta} = -W_M \frac{\partial L_M}{\partial \beta} - W_F \frac{\partial L_F}{\partial \beta} - W_M \frac{\partial L_F}{\partial \beta} - \frac{\partial C_1}{\partial \beta} < 0$$

where

$$\Delta = -U_{21} \beta g''[U_{11} U_{22} (W_F^2 U_{33} + W_M^2 U_{44}) + U_{33} U_{44} (\rho^2 U_{11} + U_{22})] < 0$$

and

$$\rho = (1 + r)^{-1}.$$

Case II: Household Constrained in the Credit Market

When the household is constrained in the credit market, and the constraint is binding, the household solves (A1.1), subject to the additional constraint that $S = 0$. This can be written as

$$\max U(C_1, C_2, L_M, L_F) \tag{A1.2.a}$$

$$\text{s.t. } C_2 = Y_1 + W_M(\overline{L}_M - R_M) + W_F(\overline{L}_F - L_F - R_F) \tag{A1.2.b}$$

$$C_2 = Y_2 + \beta g(R_F). \tag{A1.2.c}$$

Substituting (A1.2.*b*) and (A1.2.*c*) into (A1.2.*a*) and assuming an interior solution and that $g(.)$ is increasing and concave ($g' > 0$, $g'' < 0$), the utility function is additively separable and increasing and concave in each argument ($U_k > 0$, $U_{kk} < 0\ \forall k$; $U_{kk} < 0\ \forall j \neq k$), the following comparative statics results are obtained from the above problem:

$$\frac{\partial C_1}{\partial \beta} = \frac{\Omega g' W_F U_{33} U_{44}}{\Psi}$$

$$\frac{\partial L_M}{\partial \beta} = \frac{\Omega g' W_F W_M U_{11} U_{44}}{\Psi}$$

$$\frac{\partial L_F}{\partial \beta} = \frac{\Omega g' W_F^2 U_{11} U_{33}}{\Psi}$$

$$\frac{\partial R_F}{\partial \beta} = \frac{\Omega g'[U_{11}(W_F^2 U_{33} + W_M^2 U_{44}) + U_{33}U_{44}]}{-\Psi}$$

$$\frac{\partial(L_F + R_F)}{\partial \beta} = \frac{\partial L_F}{\partial \beta} + \frac{\partial R_F}{\partial \beta} = \frac{\Omega g'[U_{44}(W_M^2 U_{11} + U_{33})]}{-\Psi}$$

where $\Omega = U_2 + \beta U_{22}$ and $\Psi < 0$ is the determinant of the Hessian matrix generated from the problem. From the above result it can be concluded that:

$$\operatorname{sgn}\left[\frac{\partial(L_F + R_F)}{\partial \beta}\right] = \operatorname{sgn}\left(\frac{\partial R_F}{\partial \beta}\right) = -\operatorname{sgn}\left(\frac{\partial C_1}{\partial \beta}\right)$$

$$= -\operatorname{sgn}\left(\frac{\partial L_M}{\partial \beta}\right) = -\operatorname{sgn}\left(\frac{\partial L_F}{\partial \beta}\right) = \operatorname{sgn}(\Omega).$$

Appendix 2. Potential Inconsistency in Estimate of Gender Shock ($\hat{\delta}$) When Fertility Responds to Sex of Existing Children

Consider the following empirical model, which is (9*c*) under the assumption, made for expository convenience, that $L = 1$, and suppressing the subscript *k*, set of controls in *X*, and the household and village $\times$ time fixed effects.

$$DAYSWORK_{it} = \beta_0 Boy_{it-0} + \gamma_0 Girl_{it-0} + \beta_1 Boy_{i,t-1} + \gamma_1 Girl_{i,t-1} + \epsilon_{it}. \tag{A2.1}$$

Let

$$\begin{aligned}
&\text{(i)} \quad a \equiv \operatorname{var}(Boy) = \operatorname{var}(Girl)\\
&\text{(ii)} \quad b \equiv \operatorname{cov}(Boy_t, Girl_t)\\
&\text{(iii)} \quad c \equiv \operatorname{cov}(Boy_{t-1}, Girl_t) = \operatorname{cov}(Boy_{t-1}, Boy_t)\\
&\text{(iv)} \quad d \equiv \operatorname{cov}(Girl_{t-1}, Girl_t) = \operatorname{cov}(Girl_{t-1}, Boy_t)\\
&\text{(v)} \quad e \equiv \operatorname{cov}(Boy_t, \epsilon_t) = \operatorname{cov}(Girl_t, \epsilon_t)\\
&\text{(vi)} \quad f \equiv \operatorname{cov}(Boy_{t-1}, \epsilon_t) = \operatorname{cov}(Girl_{t-1}, \epsilon_t).
\end{aligned} \tag{A2.2}$$

(i) is true when probability of the birth of a boy equals the probability of the birth of a girl (which is true, approximately) (iii) and (iv) mean that the probability of the birth of a boy vs. a girl in the later period is independent of the sex of a child born in the first period. '*c*' and '*d*' are not equal, because a child is more likely to be born in the second period if a girl is born in the first period rather than a boy. (v) and (vi) refer to correlation of the error term with the regressors, the assumption being that while fertility itself (either in the current year or the previous year) may be correlated with the error term, the child's sex is not.

The plim's of the coefficients can be expressed as:

$$\operatorname{plim}\begin{bmatrix}\hat{\beta}_0\\ \hat{\gamma}_0\\ \hat{\beta}_1\\ \hat{\gamma}_1\end{bmatrix} = \begin{bmatrix}\beta_0 + (a-b)\{e[(a^2-b^2) + 2d(c-d)] + 2f(bd-ac)\}/D\\ \gamma_0 + (a-b)\{e[(a^2-b^2) + 2c(d-c)] + 2f(bc-ad)\}/D\\ \beta_1 + (a-b)^2[-e(c+d) + f(a+b)]/D\\ \gamma_1 + (a-b)^2[-e(c+d) + f(a+b)]/D\end{bmatrix} \tag{A2.3}$$

where $D = (a^2 - b^2)^2 + (a-b)[4bcd - 2a(c^2 + d^2)]$. While each of the coefficients individually is inconsistent due to the endogeneity of fertility, the difference in the inconsistencies in the estimates of β_1 and γ_1 equals zero, while the differences in the inconsistencies in the estimates of β_0 and γ_0 equals:

$$[e(c+d) + f(a+b)][2(a-b)(c-d)]/D. \tag{A2.4}$$

In general, $a \neq b$. Equation (12) will equal 0 under one of the following two circumstances: (1) when $c = d$ (i.e., when fertility does not respond to the sex of the child born in a previous period), or (2) when $e = f = 0$ (i.e., when the fertility outcomes are not correlated with the error term in the estimating equation).

Appendix 3. Summary Statistics

	Landless and small farmers	Medium and large farmers
Days worked – males	182	227
	(157)	(181)
Days worked – females	181	133
	(126)	(121)
Birth of girls	35	35
Birth of boys	38	50
Number of observations	344	331

References

Behrman, J. (1988). 'Intrahousehold allocation of nutrients in rural India: are boys favored? Do parents exhibit inequality aversion?', *Oxford Economic Papers*, vol.40, pp. 32–54.

Behrman, J. (1992). 'Intrahousehold allocation of nutrients and gender effects: a survey of structural and reduced form estimates'. In *Nutrition and Poverty* (S. Osmani, ed.), Oxford: Oxford University Press.

Behrman, J. (1998). 'Intrahousehold distribution and the family'. In *Handbook of Population and Family Economics* vol. I.A (M. Rosenzweig and O. Stark, eds.), New York: North Holland.

Behrman, J., Pollak, R. and Taubman, P. (1995). *From Parent to Child: Intrahousehold Allocations and Intergenerational Relations in the United States*, Chicago: University of Chicago Press.

Besley, T. (1995). 'Savings, credit and insurance'. In *Handbook of Development Economics*, (J. Behrman and T. N. Srinivasan, eds.), New York: North Holland.

Binswanger, H. and Rosenzweig, M. (1993). 'Wealth, weather risk and the composition and profitability of agricultural investments', ECONOMIC JOURNAL, vol. 103, pp. 56–78.

Browning, M. and Subramaniam, R. (1995). 'Gender bias in India: parental preferences, differential returns, or marriage costs?' mimeo, Yale University.

Deolalikar, A. and Rose, E. (1998). 'Gender and savings in rural India', *Journal of Population Economics*, vol. 11, pp. 453–70.

Foster, A. (1995). 'Prices, credit markets and child growth in low-income rural areas', ECONOMIC JOURNAL, vol. 105, pp. 551–70.

Garg, A. and Morduch, J. (1998*a*). 'Sibling rivalry and the gender gap: evidence from health outcomes in Ghana', *Journal of Population Economics*, vol.11, pp. 471–93.

Garg, A. and Morduch, J. (1998*b*). 'Sibling rivalry', mimeo, Harvard University.

Gersovitz, M. (1988). 'Saving and development'. In *Handbook of Development Economics* (H. Chenery and T. N. Srinivasan, eds.). New York: North Holland.

Jacoby H. and Skoufias, E. (1997). 'Risk, financial markets, and human capital in a developing country', *Review of Economic Studies*, vol. 64, pp. 311–35.

Kochar, A. (1999). 'Smoothing consumption by smoothing income: hours-of-work responses to idiosyncratic agricultural shocks in rural India', *Review of Economics and Statistics*, vol. 81, pp. 50–61.

Morduch, J. (1990). 'Risk, production and savings: evidence from Indian households', mimeo, Harvard University.

Morduch, J. (1995). 'Income smoothing and consumption smoothing', *Journal of Economic Perspectives*, vol. 9, pp. 103–14.

Parish, W. and Willis, R. (1993). 'Daughters' education and family budgets: Taiwan experiences', *Journal of Human Resources*, vol.28, pp. 863–98.

Rose, E. (1998*a*). 'Ex ante and ex post labor supply responses to risk in a low-income area', mimeo, University of Washington.

Rose, E. (1999*b*). 'Consumption smoothing and excess female mortality in rural India', *Review of Economics and Statistics*, vol. 81, pp. 41–9.

Rosenzweig, M. and Schultz, T. P. (1983). 'Market opportunities, genetic endowments and intrafamily resource distribution: child survival in rural India', *American Economic Review*, vol. 72, pp. 803–15.

Rosenzweig, M. and Wolpin, K. (1993). 'Credit market constraints, consumption smoothing, and the accumulation of durable production assets in low-income countries: investment in bullocks in India', *Journal of Political Economy*, vol.101, pp. 223–44.

Sen, A. (1990). 'More than 100 million women are missing,' *New York Review of Books*, 20 December, pp. 61–6.

Strauss, J. (1986). 'The theory and comparative statics on agricultural household models: a general approach', In *Agricultural Household Models: Extensions, Applications and Policy*, (I. Singh, L. Squire and J. Strauss, eds.), Baltimore: Johns Hopkins University Press.

Thomas, D. (1990). 'Intrahousehold resource allocation: an inferential approach', *Journal of Human Resources*, vol. 25, pp. 635–64.