Market Opportunities, Genetic Endowments, and Intrafamily Resource Distribution: Child Survival in Rural India

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The allocation of resources within the family has received much attention by economists in recent years. An important rationale for attempting to understand the determinants of the intrahousehold distribution of goods and time is that such behavior may importantly influence the outcomes of market-oriented policy interventions. For example, the effect of a subvention of educational institutions on the earnings potential of children may depend in part on how parents respond, by reducing or increasing the levels of household "educational" resources devoted to children. Similarly, attempts to equalize the earnings (or earnings opportunities) of men and women may result in reinforcing or counteracting substitution responses on the part of parents with respect to the distribution of household consumption or investment goods to children according to sex; for example, the time of female children may be substituted for that of the mother in household work through reductions in school attendance when wage levels increase for an adult woman (see Rosenzweig). Another important question related to the intrafamily distribution of resources is whether parents act so as to reduce or augment the effect of differences in genetic endowment on earnings or other measures of achievement (see Zvi Griliches). Unfortunately, the measurement of these potentially important phenomena appears to pose severe data requirements, necessitating either extensive household time budget surveys (see Robert Evenson; W. P. Butz and J. DaVanzo) or the imposition of a large number of unverifiable behavioral assumptions to derive inferences from more conventional data (see Edward Lazear and Robert Michael).

In this paper we examine how intrafamily resource allocations respond to changes in economic conditions and to genetic differences in children by estimating the determinants of variations in the sex-specific survival differentials of rural Indian children, based on standard census and household survey data. A number of testable implications concerning household allocation behavior are obtained from the theoretical framework without the imposition of special behavioral assumptions or restrictions on the biological determinants of child survival.

While sex differences in child survival or mortality have been examined by social scientists (see Yoram Ben-Porath and Finis Welch; R. H. Cassen; M. A. El-Badry; E. I. Hammoud; San T. Ram Gupta), we know of only two attempts to explore empirically the possibility that economic behavior may be linked to the greater mortality rates of girls relative to boys in countries such as India and Pakistan, contrary to the experience in most other places of the world, and to account for the large cross-sectional variations in sex differences in child survival rates (see P. K. Bardhan; E. Boserup). We explore in particular the hypothesis that such differences are related to the relative returns to survival, with households selectively allocating resources to children in response to variation in sex differences in their expected earnings opportunities as adults.

In Section I, we formulate a simple model of the household in which the relationships between the allocations of (unobserved) household resources between children, biological differences by sex in the responsiveness of survival propensities to consumption

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levels, adult sex-specific earnings potential, and differences in survival rates are set out explicitly. We then derive predictions regarding how changes in differential adult employment opportunities impact on household resource allocations and on sex differences in survival rates.

In Section II we specify the empirical framework to be applied to both 1961 aggregate district-level census data and 1971 household survey data from rural India. Results from the analysis of household level sex-specific survival rates are reported in Section III, and a measure of sex differences in child survival based on census age distribution information is derived and analyzed at the level of districts from all but one state in India in Section IV. Both levels of analysis indicate that the intrahousehold allocation of resources is highly responsive to market signals. In particular, it is found, consistent with the model, that where women's expected employment in the labor market is relatively high, female children evidently receive a larger share of household resources relative to male children. The sex differences in the expected productive roles of adults appear, moreover, to dominate differences in wealth or educational levels in accounting for the variation in the sex-specific survival rates of Indian children.

I. Theoretical Framework

To examine the relationship between the intrahousehold allocation of resources and "genetic" (sex) differences in potential earnings and in survival rates, and to derive testable implications with respect to the determinants of the latter, we construct a one-period household model with two types of children (male and female). It is assumed that there are direct consumption benefits as well as pecuniary contributions or transfers associated with surviving children. The household has a utility function, given by

$$U=U(x_H,m,f), \tag{1}$$

where x_H is a jointly consumed aggregate consumption good, and m and f are the number of surviving male and female children, which are treated as continuous variables. Given exogenously fixed (nonzero) levels of male and female births M and F, the household can allocate resources X_i, $i=m,f$ to the children to influence their survival.[1] The relationship between the X_i and the number of surviving children is given by sex-specific, biologically determined depreciation functions $\delta_i(X_i)$, such that

$$m=M(1-\delta_m(X_m)), \qquad \delta_m'<0; \tag{2}$$

$$f=F(1-\delta_f(X_f)), \qquad \delta_f'<0. \tag{3}$$

Let the price index of the goods used to augment survival (health) or reduce depreciation be p and surviving male and female children contribute R_m and R_f, respectively, to family resources, through direct labor contributions or transfers. The income constraint of the household ignoring discounting is

$$V+R_m(1-\delta_m(X_m))+R_f(1-\delta_f(X_f)) - X_H-p(X_m+X_f)=0, \tag{4}$$

where V is exogenous earnings or income, and M and F are arbitrarily set equal to one, so that m and f are survival rates. Note that R_m or R_f may be negative. For example, it is customary in India for the family to make a lump sum payment (dowry) to the family of the husband when female offspring marry.

The household allocates resources X_m, X_f among the children and consumes X_H to maximize (1), subject to (2), (3) and (4). The first-order conditions are

$$U_m/\lambda=-p/\delta_m'-R_m, \tag{5}$$

$$U_f/\lambda=-p/\delta_f'-R_f, \tag{6}$$

$$U_{x_H}/\lambda=1, \tag{7}$$

where λ is the Lagrangean multiplier. Expressions (5) and (6) indicate that the shadow

[1] Fertility could be endogenous without affecting our results. We do rely on the fact that the sex ratio at birth, M/F, is random.

price of the sex-specific survival good X_i, for given p, is lower the greater is the sex-specific marginal product of such goods in attenuating mortality and the higher (more positive) are the sex-specific pecuniary contributions of the surviving child of sex i.

While we are interested in how the household goods X_i are allocated between the two types of children (to M and F), assume that only the survival outcome m and f are observed, and that, as indicated in (2) and (3), survival rates are in part determined by unobserved biological processes. If we define the difference in *levels* of sex-specific survival as S, then

$$(8) \qquad S = m - f = \delta_f(X_f) - \delta_m(X_m),$$

and it can be seen that the sign of the differential mortality levels cannot indicate unambiguously the relative *levels* of the household goods allocated among the children, since the magnitudes of the sex-specific depreciation rates δ_i are also unknown. However, *changes* in the relative amounts of the X_i can map directly into differences in S, as

$$(9) \qquad dS = \delta_f' \, dX_f - \delta_m' \, dX_m.$$

Expression (9) indicates that some qualitative predictions derived from the model regarding the determinants of variations in S pertain as well to changes in the unobserved distribution of goods within the household, regardless of quantitative differences in the biological responses of male and female survival rates to alterations in the level of resources (the δ_i functions).

To obtain the effect of a change in the expected pecuniary contribution R_i on S, and thus on the relative levels of the X_i, totally differentiate expressions (4) through (7). Using expression (9) and letting own- and cross-compensated price effects for m and f be given by θ_{ii}, $i = m, f$, and θ_{mf}, and income effects by η_i, it can be readily shown that

$$(10) \quad dS/dR_f = \delta_m'\theta_{mf} - \delta_f'\theta_{ff} + f(\eta_m - \eta_f)$$
$$= \delta_m'\theta_{mf} - \delta_f'\theta_{ff} + f(dS/dV);$$

$$(11) \quad dS/dR_m = \delta_m'\theta_{mm} - \delta_f'\theta_{mf} + m(\eta_m - \eta_f)$$
$$= \delta_m'\theta_{mm} - \delta_f'\theta_{mf} + m(dS/dV).$$

If it is assumed that surviving male and female children are substitutes in (1), $\theta_{mf} > 0$, and that differences in income effects on surviving children are small, then it can be seen that expressions (10) and (11) are of opposite sign—a rise in the potential contribution (a decline in negative transfers) associated with surviving girls leads to an increase in their survival rates (consumption of good X) relative to those of boys (decreases S); a rise in R_m similarly increases the survival rates of boys relative to girls (increases S). If, however, surviving girls are a "luxury," as the evidence reported below weakly suggests, it is possible that S will decrease in response to a rise in R_m; S and R_f, however, will a fortiori be negatively related.

Thus with relatively weak assumptions imposed on the utility function and with *no* assumptions regarding sex differences in biological propensities to survive as a function of consumption levels, it is possible to derive predictions with respect to how sex differences in child survival rates, and thus implicitly the intrahousehold allocation of resources, will behave in response to changes in the economic environment. In this case, the model suggests that parental resource allocation behavior will *reinforce* market forces which operate to change sex differences in the net pecuniary contributions of offspring to the household, as long as surviving male and female children are not strong complements in the utility function.

The relative ease with which predictions regarding the association between relative survival rates and sex-specific opportunities can be derived, in the absence of information on biological determinants of survival, is due in large part to the ability to distinguish shadow price components of the family distributional commodities, X_i, which uniquely pertain to the survival of each sex. Conversely, to obtain a prediction for the effect of a change in the direct price (p) of the goods used to decrease mortality on S due, for example, to the governmental provision of health care or to increases in education levels, requires more restrictions:

$$(12)\quad dS/dp=\delta_f'\theta_{ff}-\delta_m'\theta_{mm}+(\delta_f'-\delta_m')\theta_{mf}+m\eta_m-f\eta_f.$$

It can be seen from (12) that the effect of a rise in the cost of goods used to increase child survival on *relative* sex-specific survival rates depends on biological differences in the depreciation functions, on differences in own-substitution effects and on differences in income effects. However, since the latter can be directly measured, discrepencies *in sign* between dS/dp and $dS/dV\ (=\eta_m-\eta_f)$ provide some indication of the importance of any biological difference in the sex-specific survival functions. Such tests are reported below.

II. Econometric Framework

In a stable, slowly developing society such as in rural India, parents can reasonably expect that conditions which they face as adults will also condition in a similar way the behavior of their offspring. To apply the model, we assume that the future sex-specific earnings opportunities of children as adults (E_m and E_f) affect positively their net pecuniary contributions (R_m and R_f), to the family, and that expectations of these opportunities are formed on the basis of contemporaneous sex-specific patterns of adult behavior.[2] It is these expectations that thus inform the intrafamily allocation of resources. The basic set of equations we estimate, based on the assumption of intertemporal stability, are

$$(13)\quad E_m=\alpha_{m0}+\alpha_{m1}X_1+\alpha_{m2}X_2+\alpha_{m3}X_3+\alpha_{m4}X_4+\varepsilon_m,$$

$$(14)\quad E_f=\alpha_{f0}+\alpha_{f1}X_1+\alpha_{f2}X_2+\alpha_{f3}X_3+\alpha_{f4}X_4+\varepsilon_f,$$

$$(15)\quad S=\beta_0+\beta_1E_m+\beta_2E_f+\beta_3X_3+\beta_4X_3+\beta_5X_4+\mu,$$

where X_1 is a vector of variables influencing the demand for adult labor services; X_2 is a vector of variables which may act to constrain adult employment—religious proscriptions, other segmentation; X_3 is a vector of wealth, production, or asset variables; and X_4 is a vector of variables representing educational attainment. All the X_j variables may influence the earnings of male and female adults differentially.

The error terms in equations (13), (14), and (15) are likely to be correlated as adult employment (one component of X_H in (1)) and child survival are jointly determined out of a utility-maximization process and thus reflect the same household preference orderings. Moreover, differences in the relative population sizes of adult males and females, aggregations of past intrafamily allocation decisions by parents (i.e., past mortality differentials by sex), may influence the relative returns of the two groups in the labor market, given imperfect substitution and cost of migration. To avoid simultaneous equation bias, equation (15) is therefore estimated using two-stage least squares based on the parameter estimates obtained from (13) and (14). The specification assumes that the labor-demand variables in X_1 only influence survival differentials through their effects on the expected earnings and thus on the R_f and R_m variables; all other variables jointly influence E_m, E_f, and S.

[2]Boserup notes that "In regions where women do most of the agricultural work, it is the bridegroom who must pay bridewealth, but where women are less actively engaged in agriculture, marriage payments come usually from the girl's family" (p. 48). F. L. Pryor (Table B-3, p. 364), in testing his hypotheses regarding the determination of the brideprice (minus dowery) across sixty anthropologically studied populations, accounted for 41 percent of the variance with three trichotomous variables: 1) relative time of spouses spent on economic activities (the more women work relative to men, the higher the brideprice); 2) the postmarital residence of the couple (if the couples lives with the husband's parents the brideprice is higher and vice versa), and 3) whether the husband (or wife) plays the predominant role in important family decision making.

Our measure of E_m and E_f, which is assumed to reflect the expected net economic contributions of male and female offspring, is based on adult sex-specific employment probabilities either on the household's own land, in cottage industries, or in the labor market. We use the data on employment rather than wage rates for two reasons. Wage rates may not accurately reflect the shadow value of time because exogenous cultural factors (for example, religion and caste) may prevent wives from equalizing market and household marginal products. We test for this below. Also, agricultural wage rates are not available for many districts of India, preventing comparisons between the analyses we perform on district level and household level data. Of course, female employment may vary across communities due to differences in labor market demand and household supply factors. Our estimation strategy attempts to isolate those factors which tend to affect earnings opportunities through altering labor demand or directly constraining labor supply. The sex differences in employment rates and the sex differences in market wage rates from the data below imply that the expected earnings power of males is twice that of females in rural India.

Based on the model, we would expect that $\beta_2 < 0$, $\beta_1 > 0$, with the size of the β_4 coefficients reflecting the differential income effects on m and f. If it is assumed that schooling contributes to the efficiency with which given resources are combined to augment child survival, that is, p and schooling are negatively correlated, we would expect that the β_4 and β_5 coefficients would display the same signs, as long as biological influences in sex-specific survival functions are not important (equation (12)). The signs and magnitudes of the coefficient vectors α_{i2} through α_{i4}, $i = m, f$ in (13) and (14) will reflect the influence of both demand and supply effects on adult earnings prospects.

III. Empirical Application: Rural Indian Household Data

We first use child survival data from a national sample of 4,000 rural households from India, collected and coded by the National Council of Applied Economic Research. Households were selected for analysis in which the mother was aged 20 to 55 and had borne at least one girl and one boy in districts for which wage data were available. The final sample size was 1,331 households. We test for the potential bias inherent in this sample selection in Section IV below, where aggregate data are considered for almost all Indian districts. The sex ratio at birth (males to females) in this sample, based on over 5,000 births, is 1.08, comparable to the ratio of 1.09 found by K. Pakrasi and H. Halder using the 1961–62 Indian National Sample Survey; both ratios are only slightly higher than the sex ratio at birth in the United States of 1.06. Reported sample sex ratios at birth thus do not display any unusual male bias. The survival differential variable for each household was constructed from information on the number and sex of live births, M, F (means are 2.5 and 2.3), and the number and sex of children who survived m, f (means are 2.3 and 2.0). The measure of S for household k is defined $(m_k/M_k) - (f_k/F_k)$. Table 1 lists the variables used in the analysis and their sample means and standard deviations. As can be seen, the survival differential, which is standardized for the sex ratio at birth is favorable for males, on average.

Table 2 reports estimates of the first-stage equations (13) and (14). The set of demand variables, X_1, used to identify the second-stage survival differential equations, are as a group statistically significant at the 1 percent level. As expected, in heavily Moslem areas, employment rates of women, but not those of men, are significantly lower. Similar results are obtained for areas where lower caste populations dominate. However, while factories in rural areas appear to favor the employment of men, small-scale industry, where present, tends to increase the employment demand for women relative to men.

Since household behavior can only negligibly affect district wages of men and women, the coefficients on these wage variables can be interpreted as own wage and cross wage effects on the household's demand for member's nonmarket time (see Orley Ashenfelter and James Heckman). In accord with empirical evidence from many countries, the wife's employment response to the male wage is

TABLE 1—VARIABLE MEANS, STANDARD DEVIATIONS AND LEVEL OF OBSERVATION, RURAL INDIAN HOUSEHOLDS, 1971

Variable	Mean	Standard Deviation	Level of Observation
Endogenous			
Male-Female Child Survival Differential	.0181	.281	Couple
Female Employment Rate	.479	.500	Couple
Male Employment Rate	.989	.105	Couple
Exogenous Included			
Nonearned Income (Rupees/Year)[a]	123.9	526.1	Couple
Gross Cropped Area (Acres)[a]	7.81	11.1	Couple
No Land (Dummy)[a]	.327	.469	Couple
Village Electrified (Dummy)[a]	.370	.483	Village
Normal District Rainfall (*mm*/Year)[a]	453.2	899.8	District
Wife with some Formal Education, but Less than Matriculate (Dummy)[b]	.122	.327	Couple
Head with some Formal Education, but Less than Matriculate (Dummy)[b]	.564	.496	Couple
Female Matriculate or above (Dummy)[b]	.0229	.150	Couple
Male Matriculate or above (Dummy)[b]	.194	.396	Couple
Age of Wife	39.2	8.72	Couple
Age of Husband	46.0	9.75	Couple
Exogenous Excluded			
Factory in Village (Dummy)[c]	.0865	.281	Village
Small Scale Industry in Village (Dummy)[c]	.0953	.294	Village
District Female Agricultural Wage (Rupees/Day)[c]	1.95	.896	District
District Male Agricultural Wage (Rupees/Day)[c]	3.22	1.47	District
Proportion of District Females Moslem, 15–59[d]	.0723	.109	District
Proportion of District Population in Scheduled Castes[d]	.144	.0820	District
Number of Households in Sample	1331		

[a]Variables treated as X_3 including *wealth*, production and assets.
[b]Variables treated as X_4 representing *educational* attainment.
[c]Variables treated as X_1 that only affect the derived *demand* for adult labor.
[d]Variables treated as X_2 that *culturally* constrain employment.

negative and substantial in magnitude, the elasticity being $-.45$. The wife's uncompensated own-wage employment response is also negative, but smaller in elasticity terms, $-.12$. The husband's uncompensated own-wage employment response is essentially zero, whereas his cross-wage elasticity is positive, but small, .03.

Of the other coefficients, with the exception of rainfall, all variables associated with either wealth or income tend to lower the employment probability of women, with income effects tending to reduce the employment of men somewhat less than those of women. Consistent with the hypothesis that women are employed more in wet agriculture, however, there is a significant positive correlation between normal district-level rainfall and the probability that a woman is employed; the correlation is negative for males, but not statistically significant. Schooling evidently plays an important role in determining the employment rate of women, with female schooling effects being nonlinear—women with a primary school education tend to be employed less than women with less than a primary education or with a matriculate degree or higher. Higher levels of male schooling are associated with both lower employment rates for women and, perhaps, higher rates for men. One should recall that our model has no implications for the effects of schooling on employment propensities of men and women.

The second-stage regression results for the child survival differential are reported in Table 3. The employment rate coefficients (β_1 and β_2 in equation (15)) are consistent with

TABLE 2—ORDINARY LEAST SQUARES REGRESSIONS: EMPLOYMENT RATES, RURAL INDIAN HOUSEHOLDS, 1971

Variable	Employment Rates[a] Female	Male
Nonearned Income ($\times 10^{-4}$)	−.416 (1.69)	−.214 (3.67)
Gross Cropped Area ($\times 10^{-4}$)	−33.8 (2.61)	−4.57 (1.49)
No Land	.0312 (1.00)	−.0189 (2.57)
Electrification ($\times 10^{-2}$)	−1.11 (0.37)	−.562 (0.81)
Rainfall ($\times 10^{-4}$)	.475 (2.43)	−.0552 (1.20)
Female Primary Education	−.151 (3.59)	−.0047 (0.47)
Male Primary Education	−.162 (5.20)	.0111 (1.50)
Female Matriculate	.199 (2.23)	−.0137 (0.65)
Male Matriculate	−.347 (8.10)	.0102 (1.01)
Agricultural District Wage Rate		
Female	−.0286 (1.84)	.0102 (2.78)
Male	−.0671 (7.00)	.00219 (0.86)
Presence of Factory	−.156 (3.33)	.0153 (1.38)
Small Scale Industry	.0969 (2.17)	−.0276 (2.62)
Proportion District Moslem ($\times 10^{-2}$)	−.781 (5.88)	−.0190 (0.61)
Proportion Scheduled Castes	−.0127 (7.42)	−.0000 (0.11)
Age of Wife ($\times 10^{-2}$)	−.0047 (0.02)	−.030 (0.44)
Age of Husband ($\times 10^{-2}$)	−.109 (0.42)	−.099 (1.62)
Constant	1.22 (15.96)	1.03 (57.12)
R^2	.235	.046

[a] t-statistics are shown in parentheses.

TABLE 3—TWO-STAGE LEAST SQUARES REGRESSIONS; MALE-FEMALE SURVIVAL DIFFERENTIAL, RURAL INDIAN HOUSEHOLDS, 1971

Variable	Coefficient	t
Female Employment Rate[a]	−.102	(1.95)
Male Employment Rate[a]	.317	(0.46)
Nonearned Income ($\times 10^{-4}$)	−.0310	(0.14)
No Land	.0525	(2.20)
Gross Cropped Area ($\times 10^{-4}$)	−1.23	(0.13)
Electrification ($\times 10^{-2}$)	−.818	(0.43)
Rainfall ($\times 10^{-4}$)	−.154	(1.46)
Female Primary Education	−.0084	(0.30)
Male Primary Education	−.0356	(1.57)
Female Matriculate	.0225	(0.38)
Male Matriculate	−.0753	(2.21)
Age of Wife ($\times 10^{-2}$)	.120	(0.64)
Age of Head of Household ($\times 10^{-2}$)	.105	(0.57)
Constant	−.311	(0.42)

[a] Endogenous variable.

the hypothesis that intrafamily resource allocations reinforce market signals—a rise in the expected adult male employment rate exacerbates the survival differential in favor of boys, other things equal ($\beta_1 > 0$), while the female employment rate reduces the differential ($\beta_2 < 0$), as predicted by the model. The point estimates indicate that a rise in the adult female employment rate to .66, an increase of 37 percent, would erase the mean survival differential. As indicated, however, parity in survival rates does not necessarily reflect parity in the distribution of household goods, given possible biological differences.

Holding constant the relative expected employment rates, we find evidence that increases in wealth, in terms of land ownership or other productive capital and asset income, are associated with relatively greater survival prospects of female children. As a consequence of the apparent "superiority" of female survival, boys appear to have significantly higher survival rates relative to girls in landless than in landed households.[3] The

[3] This may also reflect the fact that households with land tend to be better equipped to employ efficiently their female offspring within the household than are landless households. Women in landless households tend to work more frequently outside of the household, and this arrangement may be circumscribed by custom and religious practice, thus reducing the realized net productive contribution of women in the landless households. Note our measure of employment includes both on and off farm economic activity of women.

negative signs for three of the four schooling variables (the fourth is not statistically significant) thus suggest that biological differences in survival propensities between the sexes may be relatively unimportant, as education and income effects appear qualitatively similar.

IV. Empirical Application: Indian District-Level Data

While aggregate and regional (state) data from India on child and infant mortality rates also document significant variations and anomalies in sex differences in child survival (see Bardhan), it is well known that vital rate data from India are of insufficient quality to withstand an intensive analysis of these phenomena at a finer geographical level. Indian census age distribution data, however, appear to capture accurately sex-differences in the population by age (see P. M. Visaria). In this section we show that such data can be used to depict sex-specific differences in survival rates. Based on the constructed survival measure, we apply the model to analyze the determinants of the variation in sex-specific child mortality differences across 295 Indian districts (those which report rainfall data from the 1961 Census). These districts contain more than 95 percent of the total rural Indian population.

To derive the relationship between sex-specific age structure and mortality differentials, it is convenient to assume that the population is stable (see A. J. Coale). The sex ratio of the number of males to females aged a in a stable population is

$$(16) \qquad \frac{\bar{C}_m(a)}{\bar{C}_f(a)} = \frac{b_m e^{-ra} p_m(a) \bar{P}_m}{b_f e^{-ra} p_f(a) \bar{P}_f},$$

where b_i is the birth rate, r is the annual rate of increase, $p_i(a)$ is the proportion surviving from birth to age a, and $\bar{P}_i$ is the total number of persons of the ith sex in the population and $\bar{C}_i(a) = c_i(a)\bar{P}_i$, and $c(a)$ is the proportion age a.

Rearranging the right hand side of equation (16), we obtain

$$(17) \qquad \frac{\bar{C}_m(a)}{\bar{C}_f(a)} = \frac{b_m \bar{p}_m}{b_f \bar{p}_f} \frac{p_m(a)}{p_f(a)},$$

where the first term is the sex ratio at birth, and the second term is the sex ratio of survival rates. Constructing a ratio of the sex composition of two adjoining age groups t years apart yields

$$(18) \quad R(a+t,a) = \frac{\bar{C}_m(a+t)}{\bar{C}_f(a+t)} \Big/ \frac{\bar{C}_m(a)}{\bar{C}_f(a)}$$

$$= \frac{p_m(a+t)}{p_f(a+t)} \Big/ \frac{p_m(a)}{p_f(a)},$$

or, since $p_i(a) = e^{-\int_0^a \mu_i(x)\,dx}$,

$$R(a+t,a) = e^{\int_a^{a+t} \mu_f(x) - \mu_m(x)\,dx},$$

where $\mu_i(x)$ is the sex-specific death rate at age x. Taking natural logarithms,

$$(19) \quad \ln R(a+t,a)$$

$$= \int_a^{a+t} \left[\mu_f(x) - \mu_m(x)\right] dx,$$

it can be seen that the logarithm of the ratio of the sex ratios of adjacent age groups is the *difference* between the survival rates of females and males in the relevant age interval.[4]

[4] The closeness of the relationship between the *log* age-ratio measure and the variable of interest, the mortality differential, depends on the exogeneity of the sex ratio at birth and the lack of sex differentials in child migration rates, and the accuracy of the survey or census information on the age and sex characteristics of the population. The typical problem in using the stable population model for study of demographic processes is the strong assumption of past constancy of the fertility and mortality schedules. In this case, a change over time in fertility presents no problem unless it implies a change in the sex ratio at birth or the sex ratio of child mortality. The former appears unlikely and it is the latter which is the focus of our analysis. Moreover, while in analyses dealing with small regional populations, the stable population model is rarely appropriate because of internal migration, our attention is directed to the sex composition of children less than age 10. There is less reason to expect migration would be sufficiently sex and age selective as to distort interregional comparisons of children age 5 to 9 to those age 0 to 4. There is also the possibility that underenumeration of children differs from region to region and perhaps embodies a nonuniform error for boys and girls. Again, however, the construction of the age-ratio form of the measure of child mortality is such that sex-specific errors that persisted from one age group to the next would offset each

TABLE 4—VARIABLE MEANS AND STANDARD DEVIATIONS, INDIAN DISTRICTS, RURAL POPULATION, 1961

Variable	Mean	Standard Deviation
Endogenous		
Male-Female Child Survival Difference		
Age 0–4 and 5–9	.0391	.0446
Log Male-Female Child Survival Difference	.0415	.0471
Female Employment Rate, 15–59	.548	.257
Male Employment Rate, 15–59	.943	.0305
Exogenous Included		
Average Farm Size (Acres)[a]	12.64	13.86
Percentage Households with No Land[a]	30.61	13.98
Proportion of Land Irrigated[a]	.194	.185
Normal Rainfall (mm/year)[a]	113.9	59.09
Proportion of District Population Rural[a]	.846	.114
Proportion with Primary Education		
Females, 15–59[b]	.0259	.0349
Males, 15–59[b]	.113	.0825
Proportion Females Matriculate, 15–59[b]	.00265	.00646
Proportion Males Matriculate, 15–59[b]	.0267	.0231
Proportion of Females Moslem, 15–59[c]	.0709	.0857
Proportion of Population in Scheduled Castes[c]	.168	.098
Exogenous Excluded		
Number of Factories per Household[d]	.148	.184
Percentage of Factories with 5+ Employees[d]	4.05	4.50
Percentage of Factories Using Fuel[d]	25.12	22.37
Number of Districts	295	

[a]Variables treated as X_3 including *wealth*, production or assets.
[b]Variables treated as X_4 representing *educational* attainment.
[c]Variables treated in X_2 that *culturally* constrain employment.
[d]Variables treated in X_1 that only affect the derived *demand* for adult labor.

Table 4 reports the 1961 variables used in the district-level analysis and their sample statistics. The mean child survival difference is .039, and the logarithm of this variable suggested by the stable population model is .042. According to the above derivation, boys have had a 4 percent greater survival rate (from age 2 to about 7) than did girls in 1961, somewhat higher than the 2 percent differential observed for children ever born in the 1971 household sample. The census district data also imply similar rates for male and female adult employment as those obtained from the later household survey, with female employment rates somewhat higher in the more recent period and male employment slightly lower. This is a common pattern of postwar change in sex-specific participation rates observed in low and high income countries (see J. D. Durand).

Because wage rates are only available for a limited number of districts, only the adult employment rates are analyzed at the district level. Note that it would not be appropriate to utilize district wage rates as demand variables that identify the employment equations, as was assumed at the household level, since at the aggregate level both employment rates and wages are jointly determined. The list of identifying variables also excludes religion and caste, which at the district level seem more likely to affect directly both adult labor market behavior and sex-specific child survival. This exclusion will be reconsidered below. The remaining variables are defined analogously at the district level as they were at the household level, whenever data permit parallel specification. The proportion of the

other and any systematic omission of very young or older children would not imply a problem unless it were more common for one sex. On the accuracy of the Indian Census sex/age data, see Visaria and D. Natarajan.

TABLE 5—ORDINARY LEAST SQUARES REGRESSIONS: PREDICTION EQUATIONS, EMPLOYMENT RATES, RURAL INDIAN DISTRICTS, 1961[a,b]

Variable	Female Coefficient	Male Coefficient
Mean Farm Size	.203 (1.99)	.0164 (1.56)
Percent of Households with No Land	−.0548 (0.46)	−.0186 (1.52)
Percent Irrigated Land	−.334 (4.07)	−.00439 (0.53)
Rainfall ($\times 10^{-2}$)	.0236 (0.84)	−.0001 (0.03)
Proportion District Rural	12.63 (1.15)	−.757 (0.67)
Percent Female Primary Education	−.358 (0.44)	.140 (1.67)
Percent Male Primary Education	−.112 (0.42)	−.132 (4.89)
Percent Female Matriculate	−2.52 (0.82)	−2.08 (6.58)
Percent Male Matriculate	−.218 (0.34)	−.212 (3.21)
Proportion Females Moslem	−121.2 (8.42)	−2.10 (1.42)
Proportion in Scheduled Castes	−57.39 (4.41)	−2.75 (2.07)
Number of Factories per Household	25.39 (3.72)	1.25 (1.78)
Percent Factories with 5+ Employees	−.315 (1.07)	−.135 (4.50)
Percent Factories Using Fuel	−.0452 (0.79)	−.00453 (0.76)
Constant	67.38	98.76
R^2	.488	.617

[a] t-statistics are shown in parentheses.
[b] Shown in percent.

district population living in rural areas is added, however, to represent urban influences.

The estimated prediction equations for female and male employment rates are presented in Table 5 and appear similar to those obtained from the 1971 household data. The number of factories in the district is associated with substantially greater female employment and somewhat greater male employment rates, whereas the share of large factories with five or more employees is again correlated with lower levels of male employment. The only indicator in the census of the technology of the production units is whether factories are fuel using, which appears unrelated to employment of men or women.

Moslem districts again reveal significantly lower female employment rates, as do those districts with a greater proportion of the population in scheduled castes. In the district sample, rainfall is no longer correlated with women's employment, but instead the share of irrigated land is inversely, and farm size is directly associated with women's employment. District level primary and matriculate education of men and matriculate education of women are correlated with lower male employment, reflecting perhaps both the substitution of time from labor market activities to school work at younger ages, and a wealth effect on the demand for leisure at later ages.

The two-stage least squares estimates for rural Indian districts are reported in Table 6, with and without the inclusion of the religion and caste variables. As in the micro data, higher female employment is associated with a lower male to female survival ratio, significant at the 5 percent level ($\beta_2 < 0$). As before, the male employment coefficient is positive ($\beta_1 > 0$) and not significantly different from zero. These results suggest that the micro estimates were not mainly due to the micro sample selection procedure. According to the first regression including the Moslem and caste variables, an increase in district female employment by one-half, from 55 to 83 percent, would lower the male-female survival ratio from 1.04 to unity.

In districts in which a greater proportion of the households is landless, female child survival relative to male is higher. This is the one result where individual and district level data disagree, suggesting further study of regional and individual data on land ownership and landlessness is needed. Irrigation and rainfall are only weakly associated with improved survival prospects for girls relative to boys, representing either the enhanced

TABLE 6—INSTRUMENTAL VARIABLES REGRESSION: *Log* Male-Female Child Survival Ratio; Rural Indian Districts, 1961 Census[a]

Variable	Coefficient	Coefficient
Female Employment Rate[b]	$-.152$ (1.95)	$-.111$ (4.14)
Male Employment Rate[b]	.204 (0.33)	.0161 (0.03)
Mean Farm Size ($\times 10^{-3}$)	.158 (0.64)	.103 (0.44)
Percentage with No Land ($\times 10^{-3}$)	$-.673$ (2.50)	$-.680$ (2.70)
Percentage Irrigated Land ($\times 10^{-3}$)	$-.445$ (1.45)	$-.333$ (1.75)
Rainfall ($\times 10^{-4}$)	$-.693$ (0.99)	$-.849$ (1.37)
Proportion District Rural	.0412 (1.46)	.0346 (1.39)
Proportion Female Primary Education ($\times 10^{-2}$)	$-.108$ (0.49)	$-.0476$ (0.24)
Proportion Male Primary Education ($\times 10^{-2}$)	$-.0198$ (0.21)	$-.0423$ (.050)
Proportion Female Matriculate ($\times 10^{-2}$)	$-.229$ (0.16)	$-.547$ (0.44)
Proportion Male Matriculate ($\times 10^{-2}$)	.144 (0.75)	.114 (0.63)
Proportion Females Moslem ($\times 10^{-1}$)	$-.486$ (0.51)	– –
Proportion of Population in Scheduled Castes	$-.0303$ (0.70)	– –
Constant	$-.0565$	.0976

[a] *t*-statistics are shown in parentheses.
[b] Endogenous variable.

productivity of female labor in wet agricultural crops (see Bardhan), or the effect of increased household wealth for a given farm size in regions with more ample water supplies. There is also an indication that urban influences in the district are associated with increased female survival relative to male, a pattern noted by other authors studying the expectation of life at birth (see S. H. Preston and J. A. Weed). The coefficients on the education variables are negative in sign, consistent with their representing wealth, but they are not statistically important.

A notable finding from these estimates is that districts in which larger proportion of the population is Muslim do not exhibit distinctly different sex-specific child survival rates, holding constant for predicted women's employment rates. Indeed, the religion and caste variables are not jointly significant in the first regression, at even the 50 percent level. The assumption employed in the household level analysis that religion and caste affect the intrafamily allocation of resources to boys and girls only through their effect on adult employment rates for men and women is thus supported by the district level data. In the second regression in Table 6 these variables are excluded. In this second specification, the female employment rate coefficient is smaller, but its standard error declines by two-thirds. The precision of the estimates of the effects of landless households, irrigation and rainfall is also increased substantially.

V. Conclusions

A central working assumption underlying the economic literature on household behavior is that a family utility function exists permitting study of intergenerational allocation as an orderly optimizing process. A question for study is, therefore, whether parents allocate their investments across offspring in order to complement the distribution of genetic endowments of their children or to compensate for these endowments and thereby to equalize the economic opportunities available to their children. In a low-income country, child survival may be a sensitive indicator of parent investment in children.[5] Our analysis of data from households and districts of rural India suggests that the preference for compensating investments in offspring, assumed, for example, in Gary Becker and Nigel Tomes, does not dominate behavior, but rather the random

[5] It is possible, of course, that in families where survival rates for girls are low, more resources are invested in each girl survivor. However, across societies and over time, a negative relationship is generally observed between the level of child mortality and various forms of human capital investment, such as schooling.

differences in genetic traits of children, associated with sex in this case, evoke reinforcing allocations of family resources.[6] Children who are expected to be more economically productive adults receive a larger share of family resources and have a greater propensity to survive. These results imply that attempts to equalize the earnings opportunities of men and women in the current generation therefore may reduce the dispersion of earnings *more* in future generations than in the contemporaneous period. However, the general association indicated between household wealth and improved female relative to male survival may also imply that at least greater equality in survival opportunities between children is sought as the wealth of families increases, holding constant for the market social valuation of the genetic endowment of sex.

Appendix

District-Level Data Sources: agricultural wage rates, India Directorate of Economics and Statistics (1976); normal rainfall, irrigation, farm size, India, Directorate of Economics and Statistics (1970). All other data, India, Office of the Registrar General: age distribution, Part II-A; religion, caste, schooling, Part II-B; employment, Part II-C; factories, Part IV-B.

Individual Household Data Source: described by M. T. R. Sarma et al. and available from NCEAR, New Delhi, India.

[6]The tendency for household investments to reinforce adult market productivities is also evidenced in rural Indian child schooling and employment data. District adult female wages have twice the positive effect on the school attendance rates of girls than they do of boys, and conversely twice the deterrent effect on the employment rates of girls than of boys. In contrast, adult male wages depress child schooling and employment rates of female children by more than they do male children (see Rosenzweig and Robert Evenson).

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