Rosenzweig, Mark R.

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Educational Subsidy, Agricultural Development and Fertility Change

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Box 1987, Yale Station
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EDUCATIONAL SUBSIDY, AGRICULTURAL DEVELOPMENT AND FERTILITY CHANGE

Mark R. Rosenzweig

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The notion that the cost of increasing family size depends upon the level of expenditures or investment per child (child quality), formalized in Becker and Lewis (1973) and Willis (1973), provides a rationale for the contemporaneous inter-country negative correlation between the schooling attainment of young persons and birth rates as well as the trends in these variables over time in developed countries during their demographic transition. A sufficient condition for fertility to fall and, say, schooling to rise as development proceeds in this framework is that the shadow-price constant income effect on quality per child exceed that on numbers of children.[1] Such an explanation, however, would appear to be of little value for those who hold that population growth itself impedes economic development (e.g., Coale and Hoover (1958)). From this perspective, the compensated substitution implications of the theory are of concern, whereby price interventions which impinge on family size decisions can be used to accelerate per-capita income.

The chief focus of policies aimed at reducing fertility in the absence of income growth appears to be on altering the "own" price of children through lowering information costs associated with contraceptive methdos in order to take advantage of recent innovations in birth control technology. In this paper, we examine both theoretically and empirically the natalist impact of two alternative potential policies--reductions in the price of schooling and technological innovation in the agricultural context--based on a rural household model in which (school) investments per child influence the cost of children as in the Becker-Lewis framework and in which the returns to schooling rise in a dynamic environment as a consequence of the allocative effect of education (Welch, 1970; Schultz, 1975). I show that, as a consequence of the "quantity-quality" interaction, reductions in the direct costs of schooling may <u>raise</u> fertility levels even if child schooling and the quantity of children are substitutes as conventionally defined and even if (observed) income effects are not positive.[2]

However, we also show that if schooling improves allocative skills, the interaction makes it likely that an acceleration in the flow of agricultural innovations will tend to depress fertility, even if schooling and family size are complements, with the magnitude of the effect depending on the degree of competitiveness in the rural labor market.

Household data from India which contain cross-sectional variations in the price of schooling and in which a proportion of households were exposed to a governmental program providing increased access to the continuous flow of new, high-yielding grain varieties associated with the 'green revolution' are used to obtain estimates of the natalist impact of the alternative policies. Particular attention is paid in the empirical analysis to the use of these data to simulate a 'natural' policy experiment by the merging of the household information with district-level data. In particular, attempts are made to distinguish the response of households to the impact of technological change from cross-sectional differences in behavior associated with unobserved geographical characteristics which persist over time and which, because of the behavior of policy-makers, may be correlated with the presence of a policy intervention (Guttman, 1978).

The empirical estimates confirm both the quantity-quality interaction and allocative efficiency hypotheses, indicating that farm family fertility declined and school enrollment increased in areas affected by the dissemination of novel agricultural inputs compared to other farm households. Moreover, fertility in non-farm households was higher in those areas where local schools were easily accessible despite evidence consistent with schooling and numbers of children being substitutes.

In section I the basic model is formulated in two variants which define the spectrum of developing country rural labor market assumptions -- perfect labor mobility and the absence of a market for labor services. Section II contains a description of the data and the empirical framework, while section III reports on the results obtained and their implications for policy.

I. Theoretical Framework

To bring out the essential differences between the effects of educational subvention and agricultural technical change on fertility in developing-country agriculture I formulate a simple two-period model of the farm household. The household's utility function, given by (1), is the standard quantity-quality formulation (Willis, 1973; Becker and Lewis, 1973) in which

$$(1) \qquad U = U\,(N,\ q,\ S)$$

N is numbers of children, q is the fraction of total time (=1 unit) spent by each child in school during the first (schooling) period and S represents all other commodities. The function (1) is characterized by the usual neo-classical properties. Each child spends a fraction α_1 of non-school time in period 1 and a fraction of total time α_2 in period 2 in "own" farm production. The remaining time in each period is spent in the labor market where W_n is the first period (child) wage and ω represents the market "rental" rate of human capital or schooling services (ψ q). We characterize a "market labor" economy as $1 > \alpha_j > 0$ and a "surplus labor" economy by α_1, $\alpha_2 = 1$. Both cases are considered below.

Farm income is produced according to a production function Γ in which all variable inputs except the labor of the children, L_1, are supressed and farm scale A is assumed fixed. Total household income in the first period, with I_j representing non-earnings income in period j and a unit farm output price, is:

$$(2) \qquad Y_1 = I_1 + \Gamma\,(L_1;\ A) + N\,(1 - \alpha_1)\,(1 - q)\,W_n$$

$$\text{where } L_1 = N\,\alpha_1\,(1 - q)$$

Total income in the second period is:

(3) $Y_2 = I_2 + (1 + \gamma)\ \Gamma\ (L_2;\ A) + Nq\ \psi\ (1 - \alpha_2)\ \omega$

where $L_2 = Nq\ \alpha_2$

γ represents the farm's proportionate advance in technology from period 1 to 2 which for simplicity is assumed to augment production in a Hicks-neutral way. The degree to which output is enhanced is given by the adoption function (4)

(4) $\gamma = \gamma\ (i,\ q)$

whose inputs are the exogenous flow of innovations i (technological progress) and q, the per-child level of schooling. The dynamic allocative effect of schooling is thus represented by the positive cross-partials of the γ function. The adoption function has been expressed without the quantity of children as an argument. While it would appear that given the small scale of Indian farms, increasing the number of children of given quality would have a minimal effect on the rate of adoption, the necessary critical assumption, as will be shown below, is that an increase in the expected flow of innovations does not augment the productivity of N in adoption. Schooling is also assumed for generality to augment the productivity of a given number of children in farm production and to enhance market earnings power even in the absence of technological innovation. Neither of these latter assumptions are crucial to the implications of the model.

The two-period income constraint, ignoring discounting, is thus:

(5) $F = Y_1 + Y_2 - Nq\ \Pi_q - S\Pi_s$

where Π_s is the shadow price of S and Π_q is the non-time price of schooling (books, tuition, direct transportation costs).

a. Market Labor Model

We first consider the market labor case in which the children participate in both farm production and in the labor market; there are no impediments to labor mobility between "own" and other farms or sectors within a geographical area. In this case α_j is less than one and it can be easily shown that $\partial\Gamma/\partial L_1 = W_n$ and $(1 + \gamma)\ \partial\Gamma/\partial L_2 = \omega$ for any (non-zero) values of N and q. Maximization of (1) subject to (4) and (5) yields the first-order conditions:

$$(6)\quad \frac{U_n}{\lambda} = \left[q\ (\Pi_q + W_n - \psi\omega) - W_n\right] = q\Pi_Q - W_n$$

$$(7)\quad \frac{U_q}{\lambda} = \left[N\ (\Pi_q + W_n - \psi\omega) - \frac{\delta\gamma}{\delta q}\ \Gamma\ (L_2;\ A)\right] = N\Pi_Q - \frac{\delta\gamma}{\delta q}\ \Gamma\ (L_2;\ A)$$

$$(8)\quad \frac{U_s}{\lambda} = \Pi_s$$

where λ = Langrangean multiplier.

The full shadow prices of N and q have two components--a 'common' component Π_Q, which is multiplied by the level of the other commodity and consists of the direct cost of schooling plus time costs less the rental value of a unit of schooling services, and a 'unique' component.[3] For the quantity of children, this latter is the child wage, which offsets in part the other costs of increasing $q\Pi_Q$) and which is independent of the levels of either q or N. The unique price component for q represents the innovational returns to schooling, which is, of course, dependent on both the quantity of children and the level of q.

From (6) and (7) it can be seen that because the non-time cost of schooling Π_q enters the shadow prices of both q and N, a change in this price will have direct ("own") and indirect ("cross") effects on both commodities, while a change in the effectiveness with which q raises farm output through adoption will have only an indirect price effect on the quantity of children and a direct price effect on q. More rigorously, by differentiating totally (5), (6), (7), and (8) we obtain the total effects of changing the direct cost of schooling on q and N, where ϕ is the determinant of the bordered Hessian, the ϕ_{rc} are the relevant cofactors and $\partial q/\partial F$ and $\partial N/\partial F$ are observed income elasticities:

(9) $$\frac{dq}{d\Pi_q} = \lambda N \frac{\phi_{22}}{\phi} + \lambda q \frac{\phi_{21}}{\phi} - Nq \frac{\partial N}{\partial F}$$

(10) $$\frac{dN}{d\Pi_q} = \lambda q \frac{\phi_{11}}{\phi} + \lambda N \frac{\phi_{12}}{\phi} - Nq \frac{\partial N}{\partial F}$$

Second-order conditions constrain the first terms (corresponding to the conventional compensated own price effects) in (9) and (10) to be negative ($\phi < 0$, $\phi_{ii} > 0$). The second terms, the cross price effects, are not signed. However, the dependence of the shadow price of N on q and vice versa makes the signs of ϕ_{12} and ϕ_{21} likely to be negative. It can be readily demonstrated that:

(11) $$\text{sign}\,[\phi_{12}] = \text{sign}\,[\phi^c_{12} - \lambda\, \Pi_Q\, \Pi_s^2]$$

(12) $$\text{sign}\,[\phi_{21}] = \text{sign}\,[\phi^c_{12} - \lambda\, (\Pi_Q - \frac{\partial \gamma}{\partial q} \frac{\partial \Gamma}{\partial L_2} \alpha_2\, q\, \psi)\, \Pi_s^2]$$

where ϕ^c_{12} is the cofactor from the bordered Hessian of the standard non-interactive three-commodity consumer model, and whose sign defines in the Hicks-Slutsky convention whether N and q are complements ($\phi^c_{12} > 0$) or substitutes. Since the second terms in brackets in (11) and (12) must be positive, we see that because of the interaction between q and N the cross price effects may be positive even if N and q are weak complements and must be positive if N and q are substitutes. Thus, for a schooling subsidy, (a reduction in Π_q) to be effective in reducing fertility, it is neither necessary nor sufficient for N and q to be substitutes.

It is instead necessary that the 'unique' cross price effect $\lambda N\phi_{12}/\phi$ be positive. This latter condition is not sufficient, however, even if the income effect is ignored, because a reduction in schooling costs makes it cheaper to "consume" more children of fixed schooling level; thus the effect of an educational subsidy on N may take on any sign, even if the cross price effect is positive.[4] Second-order conditions only constrain the sum of the compensated 'common' price elasticities for N and q to be negative.

The effects of an increase in the flow of technological innovations i on q and N, given by (13) and (14), are however, predictable under much weaker conditions than those pertaining to the lowering of school costs.

$$(13)\quad \frac{dq}{di} = -\lambda \frac{\delta^2\gamma}{\delta q \delta i} \Gamma\,(L_2;\, A)\, \frac{\phi_{22}}{\phi} + \frac{\delta\gamma}{\delta i} \Gamma\,(L_2;\, A)\, \frac{\delta q}{\delta F}$$

$$(14)\quad \frac{dN}{di} = -\lambda \frac{\delta^2\gamma}{\delta q \delta i} \Gamma\,(L_2;\, A)\, \frac{\phi_{12}}{\phi} + \frac{\delta\gamma}{\delta i} \Gamma\,(L_2;\, A)\, \frac{\delta N}{\delta F}$$

Given that education's contribution to output rises with the pace of technical change as implied by the dynamic allocative hypothesis, the demand for schooling unambiguously increases with i if q is non-inferior. Morever, if (11) is negative, as is likely if q and N interact, and the income effect on numbers is small (or non-positive),[5] the demand for N will decrease in response to expected rises in the flow of innovations. As was shown above, the fulfillment of these latter conditions was not sufficient for a reduction in Π_q to lower family size. Expressions (13) and (14) also suggest that the magnitudes of the effects of technical change on q or N are positively associated with scale.

The effects of a rise in the price of child time on q and N also are relatively unambiguous compared to those due to changing Π_q, as W_n is a unique component of the full shadow price of children. With N and q weak complements or substitutes, a rise in W_n raises fertility and lowers schooling, if $\delta N/\delta F$ is small, as indicated by (15) and (16):

$$\text{(15)}\quad \frac{dq}{dW_n} = \lambda N \frac{\phi_{22}}{\phi} - \lambda (1 - q) \frac{\phi_{12}}{\phi} + [L_1 - N(1 - q)] \frac{\delta N}{dF}$$

$$\text{(16)}\quad \frac{dN}{dW_n} = -\lambda(1 - q) \frac{\phi_{11}}{\phi} + \lambda N \frac{\phi_{12}}{\phi} + [L_1 - N(1 - q)] \frac{\delta N}{dF}$$

The market labor model thus indicates that as a consequence of the dependency of the price of N on q, the direction of the effect of an introduction of a school subsidy on fertility cannot be predicted while augmenting the flow of agricultural technological innovations is likely to both increase schooling and lower family size, the effects being proportional to farm size. The model also indicates that birth rates will be higher and school enrollment likely lower where child wages are high.

b. Surplus Labor Model

In a setting in which most farm family members are not participants in the wage labor market as either buyers or sellers of labor services,[6] as depicted in the "peasant" models of Sen (1966) and Mazumdar (1975), or where labor marginal product is not closely related to observed wages (Ranis and Fei (1961), the marginal product of labor services in farm production is affected directly by the levels of N and q chosen. First order conditions for N and q, given by (6a) and (7a), have however, the

$$\text{(6a)}\quad \frac{U_N}{\lambda} = \left[q\left(\Pi_q + \frac{\delta\Gamma}{\delta L_1} - (1+\gamma)\frac{\delta\gamma}{\delta L_2}\psi\right)\right] - \frac{\delta\Gamma}{\delta L_1} = q\Pi^1{}_Q - \frac{\delta\Gamma}{\delta L_1}$$

$$\text{(7a)}\quad \frac{U_q}{\lambda} = \left[N\left(\Pi_q + \frac{\delta\Gamma}{\delta L_1} - (1+\gamma)\frac{\delta\Gamma}{\delta L_2}\psi\right)\right] - \frac{\delta\gamma}{\delta L_2}\Gamma(L_2; A) = N\Pi^1{}_Q - \frac{\delta\gamma}{\delta q}\Gamma(L_2; A)$$

same structure, with common and unique price components, as in the market model. Accordingly, the effects of a change in Π_q on N and q are equally ambiguous, with the relevant expressions the same as (9) and (10) except for mutatis mutandis changes in the bordered Hessian and cofactors. Denoting these by ϕ^1 and ϕ^1_{rc}, the relationship between the sign of the cross price effects in the surplus labor model and the usual complementary-substitutability expression is given by:

$$\text{(11a)} \quad \text{sign}\,[\phi^1_{12} - \phi^1_{21}] = \text{sign}\,[\phi^c_{12} - \lambda\,(\Pi^1_Q + \frac{\delta^2 \Gamma N}{\delta L_1^2}(1-q) - (1+\delta)\frac{\delta^2\Gamma}{\delta L_2^2}\psi^2 Nq - \frac{\delta\gamma}{\delta q}\frac{\delta\Gamma}{\gamma L_2}\psi q)\,\Pi_s\,]$$

Again, the unique cross price effect is likely to be positive because of the interaction between q and family size even if N and q are complements in the Hicks-Slutsky sense.

The direction of the relationship between the rate of technical change and q and N, given by (13a) and (14a) depends, however, on the magnitude of the marginal product of labor, in contrast to the market case.

$$\text{(13a)} \quad \frac{dq}{di} = -\lambda\left[\frac{\delta^2\gamma}{\delta q\delta i}\Gamma(L_2;A) - \frac{\delta\gamma}{\delta i}\frac{\delta\Gamma}{\delta L_2}N\psi\right]\frac{\phi^1{}_{22}}{\phi^1} - \lambda\frac{\delta\gamma}{\delta i}\frac{\delta\Gamma}{\delta L_2}q\psi\frac{\phi^1{}_{21}}{\phi^1} + \frac{\delta\gamma}{\delta i}\Gamma(\;)\frac{\delta q}{\delta F}$$

$$\text{(14a)} \quad \frac{dN}{di} = -\lambda\frac{\delta\gamma}{\delta i}\frac{\delta\Gamma}{\delta L_2}q\psi\frac{\phi^1{}_{11}}{\phi^1} - \lambda\left[\frac{\delta^2\gamma}{\delta q\delta i}\Gamma(L_2;A) + \frac{\delta\gamma}{\delta i}\frac{\delta\Gamma}{\delta L_1}N\psi\right]\frac{\phi^1{}_{12}}{\phi^1} + \frac{\delta\gamma}{\delta i}\Gamma(\;)\frac{\delta N}{\delta F}$$

Here, because of the constraints (assumed) on off-farm labor supply, the increase in i directly effects the returns to (child) labor in the second period. The fertility effect (14a) thus contains a positive first term as well as a negative cross effect, leading to an ambiguous result. If the marginal product of labor is quite low, a likely situation where the wage labor market is inoperative, however, the negative terms in (13a) and (14a) become insignificant. The same results as in the market model are then obtained--i and N are negatively and i and q positively related.

It should be noted that the predicted effects of technical innovation on fertility in each of the two models were derived under ceteris paribus assumptions, holding, in particular, wage rates constant in the market model. The total effects of sustained technical progress may thus be quite different from those derived if the labor market model is relevant, as such change may alter the demand for and supply of labor and thus alter wage rates. To the extent that wages do enter into the shadow prices of N and q, as does W_n in the first model and, the wife's wage is also a component of Π_Q, (Willis, 1973; Ben-Porath, 1973), the signs of the total effects of a change in i cannot be predicted. Moreover, families not directly benefitting from the technical progress, such as households without land, may as well alter their fertility decisions in response to the wage rate (demand) consequences of technical progress. In the empirical section, attention is thus paid to the distinction between the direct ceteris paribus technical change effects on N and q and those channelled through wages. To the extent that those former affects should be most relevant to farm (landed) households if schooling augments allocation skills, an additional prediction of the analysis is that, controlling for wage effects, there should be little or no impact of agricultural technical progress on the fertility or schooling behavior of non-farmers.

II. Empirical Analysis

a. Natural and Quasi-natural Experiments

The preceding theoretical framework suggests that although the impact of agricultural development through refueling technological innovation is more likely to have an anti-natalist impact than will school improvement as a consequence of the q-N interaction, empirical analysis is ultimately required. The quantitative magnitudes and direction of the effects of these potential policy tools on birth rates were shown to depend on the unknown preference structures of the households in the case of the cost of schooling and on the competitiveness of the rural labor market with respect to innovation, characteristics which cannot be determined a priori. To estimate the effects of sustained agricultural technical change and a school subsidy on fertility (and schooling) it is necessary, however, to obtain data based on an "experiment" in which randomly selected households are impacted by one of the two variables. Comparisons can then be made of the subsequent fertility behavior of the "impacted" and 'control' households. Such experiments may have occured unintentionally as a result of political or other developments over time or across geographical units, such as the introduction and then removal of a law, in which case the events provide a "natural" experiment if documented by data.

A data set based on a national sample survey of 5115 rural households collected in India in three rounds between 1968 and 1971 by the National Council of Applied Economic Research appears to contain close approximations to such experiments. Between 1961 and 1964, the federal government of India instituted a program, the Intensive Agricultural District Program (IADP), in which one district from each Indian state was selected to receive on a continuing basis technical assistance and assured supplies of fertilizer. By 1971 households in these districts, identified in the survey data, were subject to a significantly greater flow of new techniques, chiefly those associated with the "green revolution" high yielding grain varieties, over a 7 to 10 year span compared to prior periods. Moreover, the program was expected to continue at its outset and was funded throughout the period.[7] A second relevant "experiment" contained in these data is the existence of villages with no primary schools,(5 percent), with households residing in such villages facing therefore significantly greater schooling costs.

The principal discrepency between the ideal or natural experimental data and the "quasi-natural" experiments contained in the Indian survey is that the districts selected for the IADP and the non-school villages were not necessarily chosen randomly. If the IADP districts were selected because of pre-program characteristics correlated with fertility or schooling, for example, a variable representing the presence of a household in an IADP district will reflect district-level differences in serially correlated unmeasured variables as well as the impact of the flow of new technologies.

To obtain a more precise measure of exogenous agricultural development, regression methods can be used to purge the IADP variable of some of the pre-program systematic components. A dummy variable taking on the value of one if a district was chosen for IADP was regressed against a set of district characteristics, including schooling levels, pertaining to sixty-eight of the eighty-eight districts represented in the sample survey based on pre-IADP 1961 census data. Use of the residuals from this regression, reported below, which are orthogonal to the variables chacterizing the level of development prior to the introduction of the program, should provide a less biased estimate of the impact of agricultural development in a regression explaining fertility in 1971 than the district-level IADP variable.[8]

The OLS estimates are:

$$\text{(17)}\quad \begin{aligned} IADP = {} & 1.195 + \underset{(.0054)}{.0032}\ ENRM - \underset{(.0089)}{.0103}\ ENRF + \underset{(.0079)}{.0027}\ LITM + \underset{(.0137)}{.0153}\ LITF \\ & - \underset{(.0054)}{.0108}\ LAND + \underset{(.0038)}{.0010}\ DIST + \underset{(.0042)}{.0023}\ NLAND + \underset{(.0024)}{.0024}\ IRR \\ & - \underset{(.0014)}{.0021}\ PROD + \underset{(.307)}{.906}\ FACT + \underset{(.013)}{.060}\ SCALE \qquad \bar{R}^2 = .264 \quad n = 68 \end{aligned}$$

where LAND = average landholdings (acres), DIST = the Kuznets ratio of landholding inequality, NLAND = proportion of households without land, IRR = percentage of land irrigated, PROD = rupee value of production per acre, FACT = number of factories per household, SCALE = proportion of factories employing 10 or more workers, LITM (F) = male (female) literacy rate, rural population aged 15-44, ENRM (F) = male (female) school enrollment rate (5-14). The results in (17) suggest that selection was not random with respect to levels of industrialization or schooling. Districts with large factories and characterized by greater landholding inequality, but not higher levels of agricultural productivity, by marginally higher levels of irrigation and literacy rates, and by lower average holdings of land in 1961 were evidently selected for the program.[9]

The distribution of schools across rural villages may also be non-random, with schools likely not constructed where the village-level demand for education is minimal. The presence of a school may thus reflect other village attributes influencing schooling demand, resulting in a possibly spurious relationship between this proxy for Π_q and the household's fertility and educational investment. As a first-order test of this hypothesis, a dummy variable representing the presence of a school was regressed against the complete set of village-level variables in the NCAER data, including the presence of a health center, a factory, small scale industry, and bank or credit union; electrification; village size, and distance to nearest urban center. The only variable with a statistically significant coefficient was that for the presence of a factory. Exclusion of the set of variables other than the latter resulted in no significant change in "explanatory" power. These results thus suggest that the factory variable should be included in regression equations determining fertility and schooling (at the household level) along with the school variable if the presence of a factory in the village affects either of the behavioral variables.

An additional undesirable characteristic of the quasi-natural experiments portrayed in the Indian data, based on geographical differences in the policy variables, is that while an individual household cannot have influenced IADP implementation and school presence, a correlation of "tastes" for children and schooling with these geographical variables may be present in the data as a result of selective migration. Apparent imperfections in the land market which make the purchase and sale of land very difficult (Bardhan, 1977) greatly reduce, however, the mobility of farm households, those hypothesized to be most affected by IADP. Moreover, while such considerations are less applicable to landless, wage workers, interdistrict mobility appears to be quite low in India (Rosenzweig, 1978). We might thus expect that the relationships between the village-level school variable and the fertility of non-farm households would contain the only serious bias, although it would appear that such bias should not obfuscate the sign of the "true" relationships. We present some evidence on intervillage as well as inter-district mobility below.

b. Measurement and Specification

To obtain estimates of fertility and schooling responses to the changes in the pace of technological innovation associated with IADP, flow rather than stock measures of the behavioral variables are appropriate. Because IADP began in 1964 in about 1/2 the IADP districts, use of children ever born as a measure of fertility, as in most studies based on micro data (Willis, 1973; Ben-Porath, 1973; Schultz, 1976), for example, would obscure the impact of the program, as given the relatively low singulate mean age at marriage in India (17), the cumulative fertility of women aged 24 and over in the 1971 round of the survey would mostly reflect pre-IADP marital fertility behavior (to an extent positively correlated with age). The pregnancy rosters provided in the data allow instead both the examination of measures of additions to the stock of children after the program was in place as well as the cumulative fertility of the household prior to IADP.

The birth rate variable used in the analysis is based on the total number of births of women aged 25 to 40 in 1971 born from 1968 to 1971 which is age-standardized as given by (18) in order to take into account age patterns of fecundity.

$$(18) \qquad BRAT_{ka} = \frac{\sum_{x=a-2}^{a-.5} \beta(x)_k}{\sum_{x=a-2}^{a} n(x)} \qquad\qquad a = 25 \ldots 40$$

The n (x) are the "natural" birth rates of women aged x taken from the fertility schedule constructed from ten non-contracepting populations by

Coale and Trussell (1974). The $f(x)_k$ are the actual births of married woman k aged x in each year. The birth ratio measure (BRAT) thus reflects the degreee to which a woman reduced or "controlled" her fertility relative to a biological benchmark at any age in the 2.5 year period prior to the 1971 survey.[10] For women aged 25-40 in the sample, BRAT ranges from 0 to 1.82, with approximately 60 percent of the observations at zero.[11] Because of both the natural zero bound of the birth ratio and the concentration of observations at that bound, use of BRAT as a dependent variable in a regression framework would appear to call for the use of Tobit as an estimation procedure (Tobin, 1953).

Because fertility control within marriage may also vary with age, significant differences in BRAT in 1971 associated with IADP could merely reflect differences in age patterns of control characteristic of households in districts chosen for IADP rather than changes in desired family size. To test if the control of marital fertility differed between IADP and non-IADP districts prior to the introduction of the new technologies, we use the pregnancy roster to construct a duration ratio (DRAT) measure of cum ulative fertility for the women in 1964, given by (19),

$$(19)\quad DRAT_{ka-7.5} = \frac{\sum_{x=m}^{a-7.5} f(x)_k}{\sum_{x=m}^{a-7} n(x)} \qquad a = 25\ldots 40$$

where m is age at marriage. This fertility measure is thus standardized both for the duration of marriage and age of of the woman and reflects the average level of fertility control practised from marriage to the cutoff date marking the start of IADP. The properties of this fertility variable are described in Boulier and Rosenzweig (1978). As can be seen from (18) and (19), DRAT is simply BRAT cumulated back to m from 1964.

Total cumulative fertility may also differ between women of a given age because of differences in the age at marriage as well as due to differences in marital fertility control. ARAT, given by (20), is an age-standardized

$$(20)\qquad ARAT_{a-7} = \frac{CEB_{a-7.5}}{\sum_{x=12}^{a-7} n(x)}$$

measure of total fertility for the women in 1964, differing from DRAT only because of variations in marital duration.

Just as cumulative fertility embodies the history of birth control behavior prior to IADP, measures of the schooling attainment (highest grade completed) of children will reflect in part pre-IADP school enrollment, particularly for children aged above 10. An age-standardization procedure was, therefore, employed, comparable to that used to construct the birth rate measures. In this case, the number of children aged 5-14 currently (in 1971) in school in each household was divided by the predicted number in school based on the sample average single-year enrollment rates and the 5-14 age distribution of the household. The enrollment measure is thus:

$$(21)\qquad ENR_k = \frac{\sum_{i=1}^{n_k} e(x)_{ik}}{\sum_{i=1}^{n_k} P(x)_i} \qquad x = 5 \ldots 14$$

where P (x) is the sample proportion of children aged x in school, $e(x)_{ik}$ is variable which takes on the value of one if the ith child aged x in household k is in school and n_k is the number of children 5-14 in the household.[12]

The sub-sample of households chosen for estimation purposes consists of those with spouse-present married women, aged 25-40 with at least one surviving child aged 5 to 14, residing in the 68 districts which could be matched with aggregate district data. The reduced-form equations for each of the four dependent variables (18) through (21) are given by (22):

$$\text{(22)}\quad \begin{matrix} ENR_t \\ BRAT_t \\ DRAT_{t-7} \\ ARAT_{t-7} \end{matrix} = \alpha_{0ij} + \alpha_{1i}\, EDH + \alpha_{2ij}\, EDW + \alpha_{3ij}\, L\hat{W}H + \alpha_{4ij}\, L\hat{W}W + \alpha_{5ij}\, LWC + \alpha_{6ij}\, WLTHY + \alpha_{7if}\, LAND + \alpha_{8ij}\, FACT + \alpha_{9ij}\, SCHL + \alpha_{10ij}\, RES + \alpha_{11if}\, RES \cdot LAND + E_{ij}$$

$$i = ENR_t,\ BRAT_t,\ DRAT_{t-7}$$
$$j = \text{farm, non-farm}$$
$$f = \text{farm only}$$

where $RES = IADP - \widehat{IADP}$ from (17)

These are estimated separately for households cultivating farm land (farmers) and non-cultivating wage or salary worker households (non-farmers). All the variables employed are defined in Table 1, which also provides sample means and standard deviations.[13]

Because only 50 percent of the males and 25 percent of the females in the farm households participated in the wage labor market as sellers of labor services, wages were estimated using an instrumental variables procedure in which the natural logarithim of the daily wage rates earned by male and female adult market participants were regressed against a set of personal, village-level and district characteristics. The specifications employed are given by (23)

Table 1. Variable Definitions, Means and Standard Deviations
Farm and Non-Farm Households, Women Aged 25-40

Variable	Definition	Farm Mean	Farm S.D.	Non-Farm Mean	Non-Farm s.d.
ENR	Age-standarized enrollment index[a]	1.02	1.04	1.01	1.03
BRAT	" " marital birth rate[a]	.411	.536	.400	.526
DRAT	" " cumulative marital fertility (1964)[a,b]	.739	.596	.797	.702
ARAT	" " " total " (1964)[a,b]	.454	.310	.450	.293
EDH	Schooling attainment, husband	2.74	1.50	2.69	1.54
EDW	Schooling " , wife	1.43	.984	1.66	1.22
LWH	Natural logarithim of husband's wage (rupees)	1.115	.620	1.282	.503
LWW	" " " wife's wage	.543	.504	.477	.447
LWC	" " " child wage[c]	.255	.389	.301	.289
WLTHY	Non-earnings income	91.8	369.5	34.3	169.4
LAND	Gross cropped area	14.2	14.7	0	0
SCHL	Presence of school in village(=1)	.947	.224	.951	.215
FACT	" " factory " " (=1)	.040	.195	.146	.353
SSI	" " small-scale industry in village (=1)	.447	1.70	.568	1.53
WTHR	Effect of weather on crops (=1 if no adverse effect)	.800	.400	.794	.381
SIZE	Population size of village	1923	2541	4312	5702
IADP	Presence of IADP in district (=1)	.207	.406	.220	.415
RES	Residual from IADP equation[a]	.029	.327	.027	.306
AGEW	Age of wife	32.2	5.00	32.6	4.78
AGEH	Age of husband	34.3	6.17	35.4	5.42
n		1186		350	

[a]See text.

[b]From sample restricted to women married prior to 1963. See text.

[c]From district-level data. Source: Agricultural Wages in India, 1970-71.

$$(23)\quad \begin{matrix}LWH\\LWW\end{matrix} = \beta_{0h} + \beta_{1h}ED + \beta_{2h}AGE + \beta_{3h}AGE^2 + \beta_{4h}FACT + \beta_{5h}SSI + \beta_{6h}\,WTHR + \beta_{7h}SIZE + \beta_{8h}RES + \beta_{9h}\text{ district wage} + M_h \qquad h = LWH, LWW$$

Inclusion of the RES or IADP variable in (23) provides estimates of the wage impact of the introduction of the new technologies while the village-level employment opportunity variable coefficients--SSI, FACT, SIZE--will in part reflect impediments to intervillage labor mobility. The sex-specific district wage variable will attain significance if inter-district mobility is low, as this variable will pick up district-level differences in employment opportunities not reflected in human capital attributes. A regression similar to those specified in (23), but excluding the schooling achievement variable, was also run with the daily wages of children 5-14 as the dependent variable. Because the only variable which contributed significantly to the explanatory power of this equation was the district-level agricultural child wage, the log of this latter variable was used directly in equations (22).

The theoretical framework suggests that if the imputed wage rates accurately reflect marginal values of time in farm production, as would be true if most farm households either sell or buy labor services in the market and (local competition prevails) LWC will be positively associated with birth rates and, if income effects are small, negatively correlated with ENR, from (15) and (16). Since the adult male wage reflects the expected returns to the services of children (as well as the income potential of the father) and thus enters the common shadow price of N and q, LWH must be positively associated with either or both of the schooling and fertility variables.

The household production of child "quality", which should be positively correlated with schooling if pre-school investments contribute to learning efficiency, as depicted in most human capital accumulation models (Ben-Porath, 1968; Heckman, 1976), has been hypothesized to require significant quantities of the mother's time (Mincer, 1962; Willis, 1973). This hypothesis would suggest that in the labor market model of agriculture, increases in the value of the wife's wage in addition to increasing the returns to female children may raise the common component of the shadow prices of N and q (Π_Q). LWW thus may be negatively associated with either BRAT or ENR or both. Indeed, if the wage cost effect is dominant, the wife's wage and the price of schooling SCHL should have similar (in sign) effects on q and N except that a rise in Π_q (SCHL = 0) diminishes family resources.

The wage and school price variables should play the same roles in farm as in non-farm household behavior under competitive conditions in the labor market. The model indicates, however, that RES will be positively correlated with ENR (the own price effect) and significantly associated with BRAT only for farm households, if schooling contributes to dynamic allocative efficiency. RES should have little or no direct effect on non-farm household decisions. Moreover, if the model is correct, the sign of the RES coefficient in the farm household BRAT equation will indicate the direction of the cross price effect between q and N. The interaction term (RES•FARM) is also included in (22) for the farm household specification to test if larger farms benefit most from the returns to schooling investment in a dynamic context, as suggested by the model.

Of the other variables in (22), FACT is included to control for local industrialization, which could, independently of agricultural technical change, increase the returns to schooling.[14] WLTHY reflects non-earnings income and its coefficients will thus measure pure income effects on the schooling and fertility variables. The direct effect of landholding size (LAND) on both ENR and BRAT cannot be predicted, however, in the context in which opportunities for wage earnings do not exist, as the size of landholdings is positively correlated with the value of the time of children and thus with the price of schooling as well as with full income. Where wage rates accurately reflect the value of time, however, farm size does not affect shadow prices. Finally, the parental schooling variables are portmanteau variables, included to capture levels of household production efficiency, tastes (modern attidudes?) and awareness of contraception (Michael and Willis, 1975) among other attributes.[15]

c. Results

Table 2 reports the estimated sex-specific adult (15-65) wage coefficients in which both RES and the IADP dummy variable are included, with and without the district-level wage rate. The education coefficients indicate that rates of return to schooling are on the order of 15 percent but that life-cycle wage profiles are essentially flat. The significance of the small scale industry and village size variable coefficients suggests, however, that inter-village labor mobility is not perfect, as mobility would erase such differences unless these variables reflect compensatory premia. More importantly, wage rates appear to be from 10 (males) to 22 (females) percent higher in districts exposed to the new greeen revolution technologies based on the residual measure of IADP. The characteristics of IADP districts included in the IADP equation (17), such as mean land size and distritution, do not, however, appear to account for any of the variation in male or female wage rates, as none of the IADP coefficients attain statistical significance.

The second-stage enrollment estimates are presented in Table 3 for farm and non-farm households. As is consistent with the labor market model, the child wage has a negative effect while the adult male wage is positively

Table 2. OLS Regression Coefficients: Ln Wage Equations, All Households, Males and Females, 1974.

(standard errors in parentheses)

Independent Variable	Males (1)	Males (2)	Males (3)	Males (4)	Females (1)	Females (2)	Females (3)	Females (4)
ED	.155 (.010)	.130 (.009)	.173 (.012)	.134 (.011)	.109 (.018)	.101 (.018)	.144 (.020)	.136 (.020)
AGE	.005 (.008)	.001 (.007)	.003 (.010)	.008 (.009)	-.009 (.008)	-.012 (.008)	-.020 (.010)	-.021 (.010)
AGESQ ($x10^{-3}$)	-.041 (.091)	-.057 (.082)	-.069 (.110)	-.153 (.100)	.146 (.112)	.172 (.113)	.307 (.132)	.306 (.130)
FACT	.029 (.019)	.032 (.018)	.058 (.034)	.010 (.031)	.014 (.044)	.013 (.042)	.004 (.043)	.004 (.042)
SSI	.043 (.008)	.031 (.007)	.049 (.009)	.032 (.008)	.015 (.009)	.019 (.009)	.023 (.009)	.028 (.009)
WTHR	.004 (.039)	.082 (.037)	.032 (.049)	.076 (.044)	.034 (.037)	.080 (.036)	.078 (.043)	.114 (.042)
SIZE ($x10^{-3}$)	.044 (.003)	.029 (.004)	.044 (.003)	.030 (.004)	.071 (.012)	.054 (.005)	.070 (.010)	.055 (.006)
District Wage		.158 (.011)		.172 (.012)		.233 (.031)		.200 (.035)
IADP	.030 (.036)	.036 (.034)			.034 (.039)	.041 (.037)		
RES			.096 (.054)	.165 (.048)			.188 (.058)	.218 (.057)
Constant	.609	.230	.518	.221	.359	.023	.424	.121
$\bar{R}^2$	.355	.453	.393	.516	.382	.424	.500	.526
S.E.E.	.498	.458	.516	.461	.397	.383	.384	.374

Table 3. OLS-IV Regression Coefficients: Age-Standardized School Enrollment Rates (ENR), Children 5-14, Farm and Non-Farm Households, 1971.

(standard errors in parentheses)

Independent Variable	Farm (1)	Farm (2)	Farm (3)	Farm (4)	Non-Farm (1)	Non-Farm (2)
EDH	.211 (.026)	.222 (.026)	.210 (.026)	.222 (.026)	.217 (.038)	.226 (.038)
EDW	.049 (.035)	.056 (.036)	.048 (.035)	.056 (.036)	.108 (.055)	.117 (.055)
LWH[a]	.347 (.094)	.352 (.095)	.353 (.095)	.352 (.095)	.439 (.140)	.446 (.141)
LWW[a]	-.057 (.121)	-.067 (.121)	-.062 (.121)	-.065 (.122)	-.044 (.187)	-.043 (.188)
LWC	-.160 (.094)	-.222 (.095)	-.152 (.095)	-.228 (.096)	-.018 (.018)	-.034 (.020)
WLTHY ($x10^{-3}$)	.131 (.060)	.118 (.061)	.133 (.063)	.117 (.062)	.112 (.170)	.088 (.180)
LAND ($x10^{-1}$)	.035 (.024)	.031 (.024)	.023 (.028)	.034 (.025)		
FACT	.473 (.118)	.490 (.119)	.471 (.118)	.493 (.119)	.148 (.140)	.175 (.145)
SCHL	.185 (.124)	.221 (.125)	.189 (.124)	.220 (.125)	.008 (.193)	.033 (.194)
IADP	.296 (.075)		.254 (.093)		.260 (.177)	
RES		.173 (.084)		.202 (.118)		.110 (.153)
IADP·LAND			.004 (.005)			
RES·LAND				.002 (.006)		
Constant	-.440	-.417	-.434	-.417	-.594	-.587
$\bar{R}^2$	.290	.281	.290	.281	.383	.377
S.E.E.	.880	.885	.880	.886	.885	.889

[a] Instrumental variable.

associated with enrollment in both sub-samples. The male wage effects are statistically significant at the 5 percent level in all specifications while the child wage coefficients are statistically significant (.05 level) in the specifications with the RES variable. Consistent with the price of time argument, the female wage displays a negative coefficient, again in both types of households.

Income effects appear to be strong--the WLTHY coefficients, positive and statistically significant in the farm sub-sample, indicate an income elasticity of enrollment of about .5 while the WLTHY coefficients in the non-farm sub-sample, not measured with much precision, indicate an income elasticity of about .1.

School enrollment in farm households appears to be from 19 to 22 percent lower where villages do not have a school. As expected, farm household school enrollment is also significantly higher in the districts chosen for IADP. However, when the RES measure is used in place of the IADP dummy this differential drops from 30 to 17 percent. The latter results are thus consistent with hypothesis that parents on farms perceive the rise in returns to schooling investment when exposed to increased flows of new techniques, although the relationship is due in part to a positive income effect. Moreover, while the IADP coefficient is of similar magnitude in the non-farm sub-sample compared to that in the farm sample, when the pre-IADP district characteristics are purged from the IADP variable in specification (17) the coefficient drops substantially below that in the farm household equation.[16] A pooled regression (specification (2)) with slope dummies representing farm status, indicates that the difference in the RES coefficients across household type is statistically significant at the .05 level. The interaction coefficients in specfications (3) and (4), however, while of the correct sign, do not support the hypothesis that the returns to schooling are higher on larger farms.

Of the other variables, farm size does not appear to have an important influence on school enrollment, indicating that marginal values of time may be substantially captured by the imputed wage variables. Moreover, school enrollment appears to be significantly higher in villages that are at least partly industrialized, as represented by the presence of a factory, with the effect significantly stronger in farm households.

The Tobit BRAT coefficients, estimated using maximum-likelihood, are reported in Table 4. These also appear to be consistent with the hypothesis that farm households respond much more strongly to agricultural technical change in terms of their fertility behavior than do non-farm households. Birth rates in farm households are about eleven percent lower, controlling for the wage effects of technical change, in districts where the flows of new technologies were more rapid, while no significant differences in birth rates are exhibited by the non-farm households in the IADP districts. This result appears robust to the measurement of IADP variable. The cross-price effect in (14) thus appears tobe positive in farm households and to dominate the income effect which appears from the WLTHY coefficients to be very small.[17] Moreover, the pre-IADP fertility estimates, reported in Table 5, indicate that prior to the introduction of the new

Table 4. Maximum-Likelihood Tobit Coefficients: Age-Standardized Marital Birth Rates (BRAT), 1968-1971 Farm and Non-Farm Households.

(standard errors in parentheses)

Independent Variable	Farm (1)	Farm (2)	Farm (3)	Farm (4)	Non-Farm (1)	Non-Farm (2)
EDH	.107 (.067)	.101 (.067)	.104 (.067)	.101 (.067)	-.332 (.138)	-.337 (.138)
EDW	-.013 (.137)	-.013 (.064)	-.014 (.063)	-.013 (.064)	.228 (.119)	.230 (.119)
LWH[a]	.371 (.413)	.325 (.412)	.361 (.412)	.326 (.413)	1.672 (.896)	1.708 (.896)
LWW[a]	-.660 (.476)	-.648 (.478)	-.639 (.477)	-.648 (.477)	-1.650 (.692)	-1.680 (.694)
LWC	.011 (.132)	.043 (.131)	.016 (.130)	.044 (.131)	.485 (.295)	.492 (.287)
WLTHY ($x10^{-3}$)	.100 (.140)	.100 (.141)	.100 (.144)	.100 (.141)	.300 (.483)	.300 (.469)
LAND ($x10^{-1}$)	.003 (.029)	.001 (.029)	.011 (.028)	.001 (.040)		
AGEW	-.059 (.006)	-.059 (.006)	-.059 (.006)	-.059 (.006)	-.069 (.016)	-.069 (.016)
FACT	.201 (.189)	.209 (.190)	.196 (.190)	.208 (.190)	.161 (.220)	.192 (.224)
SCHL	-.182 (.164)	-.212 (.163)	-.179 (.165)	-.211 (.164)	.099 (.081)	.102 (.079)
IADP	-.245 (.106)		-.297 (.133)		-.044 (.194)	
RES		-.224 (.110)		-.232 (.168)		-.103 (.256)
IADP·LAND			-.004 (.006)			
RES·LAND				-.001 (.008)		
Constant	1.457 (.334)	1.461 (.334)	1.476 (.335)	1.460 (.334)	1.933 (.738)	1.887 (.740)
$\bar{R}^{2}$[b]	.056	.054	.055	.053	.046	.047
S.E.E.	.522	.522	.522	.522	.512	.512

[a]Instrumental variable.

[b]From OLS regression.

Table 5. .OLS-IV Regression Coefficients: Age-Standardized Cumulative Marital (DRAT) and Total (ARAT) Fertility Prior to IADP, Farm and Non-Farm Households, 1964.

(standard errors in parentheses)

Independent Variable	Farm				Non-Farm			
	DRAT		ARAT		DRAT		ARAT	
	(1)	(2)	(1)	(2)	(1)	(2)	(1)	(2)
EDH	.010	.010	.024	.024	-.127	-.125	.029	.030
	(.037)	(.037)	(.019)	(.019)	(.067)	(.068)	(.029)	(.029)
EDW	.036	.036	-.023	-.023	.053	.050	-.009	-.008
	(.037)	(.037)	(.019)	(.019)	(.058)	(.059)	(.025)	(.025)
LWH[a]	.162	.157	.018	.019	1.237	1.181	.226	.204
	(.259)	(.259)	(.137)	(.137)	(.426)	(.427)	(.181)	(.181)
LWW[a]	-.039	-.041	.082	.082	-.701	-.672	-.121	-.116
	(.228(	(.223)	(.117)	(.118)	(.324)	(.325)	(.138)	(.138)
LWC	.046	.037	.068	.070	.100	.148	.047	.076
	(.071)	(.071)	(.038)	(.038)	(.141)	(.138)	(.060)	(.059)
WLTHY ($x10^{-3}$)	.056	.057	.019	.019	.191	.223	-.018	-.0002
	(.061)	(.061)	(.030)	(.031)	(.240)	(.241)	(.100)	(.100)
LAND	.100	.101	.180	.180				
	(.141)	(.141)	(.075)	(.075)				
AGEW	-.024	-.024	-.003	-.003	-.033	-.034	-.003	-.003
	(.004)	(.004)	(.002)	(.002)	(.009)	(.008)	(.004)	(.004)
SCHL	.015	.013	-.026	-.025	.158	.182	.031	.029
	(.096)	(.096)	(.051)	(.051)	(.186)	(.187)	(.079)	(.080)
FACT	.170	.171	.028	.028	.309	.286	.024	.029
	(.113)	(.113)	(.059)	(.060)	(.114)	(.117)	(.049)	(.050)
IADP	-.014		.002		-.178		-.104	
	(.056)		(.030)		(.109)		(.049)	
RES		.005		-.002		-.030		-.104
		(.068)		(.036)		(.136)		(.058)
Constant	1.228	1.224	.456	.457	2.088	2.074	.450	.414
$\bar{R}^2$	.049	.049	.019	.019	.097	.089	.062	.056
S.E.E.	.581	.581	.306	.306	.677	.670	.283	.284

[a]Instrumental variable.

technologies, no significant difference in either marital fertility control (DRAT) or total cumulative fertility (ARAT) existed between farm households, although total fertility may have been slightly lower among non-farm households in the districts that were ultimately chosen for the program.

In contrast to the IADP and RES estimates, none of the coefficients of the school presence variables achieve statistical significance by conventional standards in the BRAT equations and, as is consistent with the interactive quantity-quality model, where schools are present (Π_q is low), fertility rates of the non-farm households are higher. This latter result indicates the dominance of the negative own price over the positive cross price effect (given the low income elasticity of children). Moreover, the presence of factories, which appeared to increase enrollment rates, also is positively associated with birth rates in farm and non-farm households; industrialization without sustained technical change thus does not appear to necessarily lower fertility.

As in the enrollment equations, the child wage coefficients display signs in accord with the predictions of the model, being positively associated with birth rates in the two sub-samples. The negative coefficients on the wife's wage appear to additionally support the price of time hypothesis and parallel the results obtained by Rosenzweig and Evenson (1977) based on district-level data from India for 1961. As was noted, because the adult male and female wage rates enter the common shadow price of q and N, the former reducing and the latter on net augmenting Π_Q, the adult wage coefficients should echo the SCHL effects in sign patterns. Since the positive LWH and negative LWW coefficients in all equations indicate the pervasive domination of own over cross price effects, the negative sign of the school presence variable coefficient in the farm household sub-sample fertility equation represents the only result inconsistent with the structure of the model. Thus in the 5x4 matrix of adult wage, SCHL and RES coefficient signs for the two dependent variables in the two sub-samples, only one sign is "wrong" (although not statistically significant).

The estimates reported in Tables 3 and 4 thus indicate that 1) the cross price effect between the quantity of children and child quality (as represented by school enrollment) is positive, as indicated by the RES coefficients in the farm household BRAT equation and 2) that when the effects of a variable changeon ENR and BRAT are composed of own and cross price effects, the former tend to dominate, so that reductions in the price of schooling appear to have a small but positive effect on non-farm birth rates.

Computation of the total (wage and direct) effects of the IADP-related disequilibria, from the RES coefficients in Tables 2 and 4, indicate that where IADP was introduced birth rates fell by 12 percent in farm households and by 5 percent in non-farm families, the latter due almost totally to the increase in the female wage. By summing over age-specific birth rates in the two sub-samples from the mean ages at marriage to age 45, these reductions would translate into a decrease in completed family size of .72 and .32 children respectively. Aside from the qualifications already discussed, these last estimates must be interpreted with caution, however, as they are conditional on a constant marriage age, which itself may change in response to agricultural development.

III. Conclusion

In this paper we have attempted to integrate in a single model the theoretical literature on the interaction between the quantity of and investment in children with that on the allocative roles of schooling in a dynamic context to shed light on the effects of educational subvention and agricultural innovation on fertility change in a developing country context. Household data from India which contain a geographically selective agricultural development program promoting technical change to farmers and village-level differences in school accessability were used to estimate the effects of changes in the shadow prices of children and schooling investments per child on fertility and school enrollment. The empirical estimates were consistent with the model, suggesting that farm households responded to their exposure to new agricultural inputs by both increasing schooling investment and by lowering family size even though the new technology appeared to increase the demand for labor services. However, school proximity, while increasing schooling, appeared to be either negligibly or positively correlated with birth rates. These results thus indicate that the returns to investments in agricultural research and dissemination in terms of their efficacy in promoting economic development may be even higher in a country such as India than those high levels already documented (Evenson and Kislev, 1975).

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Footnotes

[1]In Becker and Tomes (1976) it is shown that if the contribution of child endowments to total child quality is taken into account, the "true" income elasticities of numbers of children and child quality may be equal and of average magnitude but economic development may depress fertility and raise investments in children. However, it is also pointed out that an increase in the rate of exogenous income growth, by augmenting endowments relative to potential income, will tend to raise family size and lower expenditures on children. In this paper we argue that accelerations in economic growth brought about by technical change also invoke price effects which are likely to cause fertility to fall as a consequence of the interaction between the quantity and quality of children.

[2]DeTray (1976) considers school subsidization as a possible policy tool aimed at reducing fertility. However, he ignores the interaction between quality and quantity in incorrectly arguing that a negative coefficient in a family size regression on a variable representing child quality, even if consistently estimated, is evidence that reductions in the price of schooling would lower fertility, as we demonstrate below.

[3]The unique prices correspond to the 'fixed' prices introduced in Becker and Lewis (1973). In this model, the latter are not necessarily exogenous.

[4]Rosenzweig and Wolpin (1978) demonstrate that the sign of the effect of an exogenous change in family size on schooling (child quality) provides the sign of the <u>unique</u> cross price effect between q and N. Similarly, the direction of the effect of an exogenous rise in schooling on N, induced by

a compulsory schooling law, for example, which would provide the sign of ϕ_{12}, would not necessarily indicate whether lowering the direct costs of schooling or increasing school accessability would depress fertility (as implied in DeTray, 1976), as indicated by expression (10).

[5]While child endowments are not explicitly considered in the model to reduce complexity, in the context of a developing country we would expect that the ratio of total to endowed child quality would be quite low. Becker and Tomes (1976) show that this implies that the observed income elasticity for numbers of children is likely to be substantially lower than that for child schooling. This implication is confirmed below.

[6]It is shown in Rosenzweig (1977) that if family and hired workers are close substitutes, as long as households either buy or sell labor, market wages will accurately depict the marginal products and value of time of working family members who do not participate in the market as sellers of labor services.

[7]See Gaikwad *et al*. (1977) for a description of the IADP program. It is important to note in the context of the allocational efficiency hypothesis that the green revolution was not a "one-shot" introduction of a new technology but rather represented the beginning of a continued flow of new grain varieties as well as new problems associated with disease resistance, fertilizer use, irrigation etc. It is the 'refueling' or continuous nature of the green revolution technologies that augment the returns to investment in education.

[8]Consider the "quasi-fixed effects" model:

$$IADP = \alpha_1\ \xi_1 + \varepsilon_1$$

$$N = \beta X + \beta_1\, IADP + \beta_2\ \xi_2 + \varepsilon_2$$

where the α and β are coefficient vectors, the X are exogenous determinants of fertility, the ξ_i are vectors of unobserved geographical characteristics and the ε_i are random error terms, with cov $(\varepsilon_1, \varepsilon_2) = 0$. The bias in β_1, the estimated effect of the program on N, arises if cov $(\xi_1, \xi_2) \neq 0$ (geographical characteristics persist over time) and both α_1 and $\beta_2 \neq 0$. The method proposed in the text to eliminate the consequent covariance between IADP and the compound error term in the fertility equation is to find a set of exogenous instruments Z correlated with ξ_1 and to estimate the equation IADP = $\gamma Z + \varepsilon_3$, where ε_3 is orthogonal to the Z and contains the original random error ε_1 plus that part of the ξ_1 term not correlated with the Z, say ε_1'. If cov$(\varepsilon^1{}_1, \varepsilon_2)$ and cov$(\varepsilon^1{}_1, \xi_2) = 0$ then the regression $N = \beta^1 X + \beta_1 \cdot (IADP - \hat{\gamma}\ Z) + \beta_2\ \xi_2 + \varepsilon_2$ will provide consistent estimates of the effects of IADP on N. For consistency, it thus is necessary that the Z reflect all of the variance in ξ_1 that is correlated with ξ_2.

[9]Estimation of (17) using maximum-likelihood logit produced similar results except for a substantial decrease in the ratio of the LAND coefficient to its asymptotic standard error. The (unbiased) OLS estimates are used in the subsequent second-stage regressions to conform to the linearity assumptions required for proofs of consistency.

[10]To see the implications of the biological constraint on fertility due to the covariance between age and fecundity, consider the following simple non-stochastic identity $f(a) = n(a) \cdot (1 - p)$, where $f(a)$ is the birth rate of a woman in age interval a, $n(a)$ is her potential fertility at age a and $(1 - p)$ is the degree to which fertility is controlled. If fertility control is a linear function of a set of X variables such that $(1 - p) = \beta X$, then the effect of a change in any X_i on f is $\frac{\partial f}{\partial X_i} = n(a) \cdot \beta_i$, which is a function of the age of the woman, since $\frac{\partial^2 f(a)}{\partial X_i \, \partial a} = \frac{\partial n(a)}{\partial a} \cdot \beta_i$. Division of actual births by an approximation to potential fertility (natural fertility) results in the equation $f(a)/N(a) = (1 - p) = BRAT$, whose derivatives are independent of age. An alternative procedure, stratification of the sample into narrow age groups, would also reduce the problem but would result in a reduction in degrees of freedom, with consequent loss of estimation precision. For a more complete discussion, see Boulier and Rosenzweig (1978).

[11]Natural fertility should not be confused with maximum fertility. Populations displaying natural fertility behavior are similar with respect to age-patterns of birth rates, but differ in fertility levels. The bias in linear estimating equations which do not take biological constraints into account arises from the non-linear age pattern of fecundity, reflected in the $n(x)$ schedule.

[12]Note that the standardization procedure employed is superior to the use of dummy variables as regressors depicting the age-composition of children since such variables will reflect, in part, family size.

[13]The extent to which actual fertility is depressed below potential fertility may vary with age, as is true in the sample population studied. An age variable was thus included as a regressor in all the fertility equations. For women in farm households in the 4 5-year age groups 25-40, birth rates are .550, .407, .272 and .152 respectively while the birth ratios (BRAT) are .499, .390, .304, .236.

[14]It will be recalled that school and factory presence were found to be positively associated.

[15]We thus refrain from discussing the estimates associated with these variables in subsequent sections. The reader may supply his own interpretations. It should be noted, however, that exclusion of the parental schooling variables from the fertility and enrollment equations does not significantly alter the results obtained, although they do contribute significantly to the explanatory power of all equations.

[16]Inclusion of the complete set of 1961 district characteristics used to estimate (17) in all the second-stage equations in addition to RES added significantly to explanatory power in all cases but did not alter the reported coefficients importantly. The parental schooling variable coefficients tended to decrease in mangitude and statistical significance, however.

[17]The relative magnitudes of the estimated income effects in the birth rate and enrollment equations are consistent with the hypothesis proposed by Becker and Tomes, given the relative importance of child endowments in child quality in a country such as India. However, it should be noted that the WLTHY coefficients reflect past savings decisions which may in part be correlated with preferences for numbers of children versus child quality.