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THE IMPACT OF GOVERNMENT INTERVENTIONS ON HEALTH, SCHOOLING AND FAMILY PLANNING IN THE PHILIPPINES

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The substitution and complementarity relationships among child health, child schooling, and number of children in a developing-country context are explored, in order to test the thesis that governmental initiatives in health care, education and family planning can be orchestrated in a manner that would permit a cost-effective implementation of the national goals of reduced population growth and increased human capital buildup. Government policy initiatives and an economic framework of household decision-making are linked together, using a reduced-form demand model which incorporates unobservable individual and community effects. An instrumental variables estimation technique is used to resolve the combined problem of unobserved effect and the endogeneity of policy variables. The estimation results suggest that households respond optimally to variations in the price of governmental programs by shifting the allocation of family resources from an assured number of children to less but healthier and better-educated children. A carefully designed policy that combines government programs in a mutually reinforcing way would be much more effective than a one-program instrument in improving household welfare.

Introduction

Social services which stand at the ready disposal of the government — health centers, schools, family planning clinics — may be considered attempts to shift resources from increasing family size (quantity) to augmenting health and education. These government programs — provided free or at subsidized cost — reduce the price of health inputs and schooling and also the price of contraceptives, through direct subsidies, or indirectly by improving access to these services. They also lower the cost of acquiring and decoding information on health care practices and contraceptive technology. As a result, low-income households are able to get around the household budget constraint and respond to differences in relative costs and prices, inducing in the

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process a substitution in the family context between the quantity of children and the quality embodied in them.

The purpose of this study is to assess the impact of these government services on human capital buildup and fertility reduction in the Philippines. The focus is on the household as the crucial link between broad public policies, local government programs and family and individual welfare (Rosenzweig, 1986). The stress on the family as the decision-making unit implies, in a broader sense, that changes in household (and societal) behavior can alleviate the disagreeable consequences of rapid population growth (Birdsall, 1988). The question is how to create public policy changes that would permit families to substitute away from children and allocate more resources to child health and education. The key lies in sorting out the household's behavioral response to existing health, schooling and family planning programs of the government, and in fashioning a cost-effective mix of program interventions that can induce substitution effects (toward greater child quality) large enough to compensate for income effects generated by the same programs and other household endeavors. The thesis is that efforts to provide publicly subsidized social services in the form of health care facilities, schools, and family planning programs can be orchestrated to implement the joint public policy goals of decreased population growth and improved human capital formation.

The methodological approach is to link together government policy initiatives and an economic model of household decision-making. The study considers a model of household choice over time and across heterogenous localities, in which child health, child schooling and number of children are the "household products" and which incorporates unobservable individual and community effects. The central coefficients are the price effects of governmental activities on health, education and family planning. The model merges household economic and demographic variables with community-level determinants (public sector programs, locational characteristics, consumer prices).

It is assumed that public provision of social services is likely to be non-random, that is, there are unmeasured time-invariant, location-specific endowments that influence the distribution of health, education and family planning infrastructure. Area-level public sector programs such as health services are likely to be targetted to communities with poor health environments and that lack adequate health care infrastructure, or with other population characteristics that the program itself is designed to change. Because of this unobserved heterogeneity, panel analysis[1] is exploited to avoid bias and inconsistency in the empirical estimation. This requires the use of an instrumental variables estimation technique proposed by Hausman and Taylor (1981) to control for unobservable effects of coefficients associated with both time-varying program variables and time-invariant (and likewise time-varying) socio-economic background and community variables.

Past Studies

Recent empirical investigations on how family planning, medical care and schooling infrastructure influence fertility, child health and child education in developing countries so far have offered mixed signals. They provide no definitive evidence of how the household choice variables of child quantity and quality are affected by government policy measures.

Rosenzweig (1982) shows, using maximum likelihood Tobit estimates, that schools are inversely related to marital birth rates, which is the a priori expectation. Paqueo (1978) finds that public education expenditure is negatively associated with fertility, but is significant only in the low education/rural subsample. Strauss (1988) parts ways with this widely-held pattern: in his finding, school distance coefficients have negligible impact on height for age measures.

Strong evidence of the positive impact of family planning on weight and height of children is found in Rosenzweig and Wolpin (1986), which also finds that family planning subsidies either complement government health infrastructure, or substitute for it in the presence of health externalities. The asociation is reversed in Horton (1986), however, and is attributed to factors linked to the subjective costs of family planning.

A good number of studies focus more on joint program effects and substitution possibilities between fertility decline and human capital investment. Strauss (1988) finds negative interactions of health and schooling facilities with maternal education, implying complementarity of child quality inputs with mother's schooling. In a similar vein, the availability of medical and family planning services is found by Rosenzweig and Schultz (1982) to decrease levels of fertility and child mortality while controlling for mother's education. That breastfeeding is

[1]The estimation method used in this study is associated with a linear static model. To use panel data in estimating dynamic behavioral relationships would involve models containing lagged dependent variables. See Hsiao (1986).

highly substitutable with many government programs designed to lower fertility is evidenced in Anderson (1984). More systematic studies of the joint effects of government programs on household choice quality and quantity variables are those of Rosenzweig and Wolpin (1982) and Hossain (1989) both of which provide compelling evidence on substitutability among family planning, health care and education facilities.

Unobserved Heterogeneity

Unobserved heterogeneity — the possible systematic correlation between compensatory public services and unobservable individual and community effects — is dealt with only in a few studies. The use of a fixed-effects model is explored in Strauss (1988) to eliminate the systematic correlation among unobserved household characteristics, and time-invariant policy variables. The drawback is that the latter vanish from the model, although not their interactions with maternal education. In a somewhat different vein, Adair, *et. al.* (1988) takes into account unobserved exogenous characteristics by employing a dynamic model which continuously substitutes out for lagged endogenous values of health variables affecting consumption of both health and non-health related factors. Using a fixed-effects model for panel estimation, Rosenzweig and Wolpin (1986) account for child-specific and location-specific health endowments.

The Household Production Model

The model is a variation of Rosenzweig and Wolpin (1982), Rosenzweig and Schultz (1982) and Anderson (1984). Becker and Lewis (1973) is used as a point of departure. To focus on parental preference orderings over both family size (quantity) and characteristics such as health, nutrition and education (quality) of their children, parents are assumed to maximize a utility function of the following form:

$$U\,(Z_j,\ j{=}N,S,E,L)$$

where

Z_N is the quantity of children,
Z_S is their health index,
Z_E is their average schooling, and
Z_L is a composite of other consumption goods representing the household's standard of living.

The variables Z_N and Z_E are components of child quality. Each Z_j is produced in the home. The production functions, which are linearly homogeneous, are of the form

$$(1)\quad Z_j = Z_j(\{X_j, t_{i,j}, F_i\},\ i = m,f;\ j = N,S,E,L) + \alpha$$

where X_j are aggregate bundles of goods, purchased at market price p_j and used in the production of each good j. $t_{i,j}$ are time inputs used by husband m and wife f to produce Z_j. Home production is subject to the efficiency of the parents in the production process (F_f and F_m), that is, in combining household resources — mainly household time and household capital — with acquired market goods to produce the final commodities that yield utility. It is assumed that the skills used in producing $Z_N(F_i)$ are identical to those used in $z_E(F_i)$ and $z_S(F_i)$. These production functions exhibit constant returns to scale and no joint production. α are exogenous, environmental endowments.

If V is the non-wage income of the household, and w_i is the wage rate, the full income constraint is defined as

$$(2)\quad Y = \Sigma w_i t_{i,G} + V = \Sigma p_j x_j$$

where $t_{i,G}$ is the time spent in the market. Market income is thus equal to spending on market goods. The opportunity cost of time spent in home production is the market wage. Rewriting market time $t_{i,G}$ in terms of full time T_i and time spent in home production, and adding all $w_i t_{i,j}$ terms to the sum of market purchases to form the shadow prices of Z_j, the full income constraint becomes

$$(3)\quad Y = V + w_i T_i = \Pi_j Z_j$$

where Π_j is the shadow price or full unobservable price of Z_j. The budget restraint, with an exogenously given income, Y, is

$$(4)\quad \begin{aligned} Y &= Z_N \Pi_N + Z_N Z_E \Pi_E + Z_N Z_L \Pi_L + (Z_{Nmax} - Z_N)\,\Pi_C \\ &= Z_N (wt_N + p_N x_N) + Z_N Z_E (wt_E + p_E x_E) \\ &\quad + Z_N Z_L (wt_L + p_L x_L) + (Z_{Nmax} - Z_N)(wt_{z_{Nmax}} - Z_N + p_c x_c) \end{aligned}$$

where p_C is the unit price of averting birth through contraceptives, and x_j and t_j are the marginal input coefficients of goods and time used in the household production of commodity j. $Z_{Nmax} - Z_N$ is the number of births averted (the difference between natural fertility and desired fertility).

In this model, the time inputs are ignored and so is the interactive cost relation between quantity and investments for each child. Removing them from the model will neither alter the reduced-form demand relationship nor add testable predictions to the demand framework. Following Rosenzweig and Wolpin (1982), the budget constraint can be simplified and rewritten as

$$(5) \quad Y = Z_N P_N + Z_E P_E + HP_H + (Z_{Nmax} - Z_N)\, P_C + Z_L P_L$$

where P_N is the gross price of a live birth, P_C is the price of preventing birth through contraceptives, $P_N - P_C$ is the net price of a live birth, which consists of delivery and prenatal expenses including the opportunity cost of mother's time net of contraceptive costs. Alternatively, it is the shadow price of fertility which is equal to the market price of children net of the market price of contraceptives. P_E, P_L, and P_H are the costs of Z_E, Z_L, and H, respectively, where H are purchased health inputs assumed to positively affect the survival of children Z_N. That is, survival is a function of health goods and unobservable exogenous health endowments:

$$Z_S = Z_S\,(H) + \alpha.$$

The problem for the household is to

maximize $\quad U\,(Z_j,\ j = N,S,E,L)$

subject to $\quad Y = Z_N P_N + Z_E P_E + HP_H + (Z_{Nmax} - Z_N)\, P_C + Z_L P_L$
$\quad Z_S = Z_S\,(H) + \alpha.$

For infinitesimal changes at equilibrium values, it is possible to decompose demand for the household goods into price and income effects on the consumption of Z_N, Z_E, Z_S, and Z_L of changes in the prices of public goods (family planning centers, health services, and schools). The decompositions are embodied in the Slutsky expressions:

$$\frac{dZ_i}{dP_C} = \frac{(dZ_i)}{(dP_N)_{dU=0}} - (Z_{Nmax} - Z_N)\,\frac{(dZ_i)}{(dY)_{prices\ const}}$$

$$\frac{dZ_i}{dP_H} = \frac{1}{MP_H}\frac{(dZ_i)}{(dP_H)_{dU=0}} - Z_s\,\frac{(dZ_i)}{(dY)_{prices\ const}}$$

$$\frac{dZ_i}{dP_E} = \frac{(dZ_i)}{(dP_E)_{dU=0}} - Z_E\,\frac{(dZ_i)}{(dY)_{prices\ const}}$$

where $i = N,S,E,L$ and MP_H is the marginal product of health goods in the production of child health and nutrition.

The Slutsky equations offer a way to make predictive statements from the symmetry of compensated price effects if a broad range of demand relationships is considered. To the extent that they suggest possibilities of substitution tradeoffs among public services as relative prices are varied, the resulting mix of household choices can be determined. Consider reducing P_C by the subsidization of a family planning program. The number of children would go down as a fall in the price of contraceptives would increase the cost of raising children (that is, $P_N - P_C$ would go up and lead to a drop in demand for Z_N). In addition, a small P_C means that if quantity and quality are held to be consumption substitutes $[(dZ_E/dP_N)_{dU=O} > 0, (dZ_S/dP_N)_{dU=0} > 0]$, the price of goods used in the production of both the education and health variables will be reduced, implying increased schooling and health of children.

In addition, if, as income increases, quality and quantity are substitutes in utility, subsidization of health inputs and/or schooling will reduce fertility — reinforcing the price effect of reduced contraceptive costs — as long as the compensated price effect outweighs the income effect.[2] If health subsidies go to prenatal care, however, the cost of producing Z_N would drop, and there would be an unambiguous rise in fertility. Finally, with small income effects, schooling and health — the two components of child quality considered here — must be considered complements (cross-price effects have the same sign). All these can be confirmed only empirically; what is known a priori is that the cross effects are equal. The direction of the relationship has to be verified. In general, the relative magnitudes of income and substitution effects determine the change in demands for children and child services.

Statistical Implementation

The econometric household model yields the following demand equation for Z_i, $i = N,S,E,L$:

$$(7) \quad Z_i = Z_i (P_N, P_C, P_H, P_L, P_E, Y, F, \alpha)$$

[2]The family would experience a rise in real income as a result of the subsidies, which will likely relax the household budget constraint, induce a decline in the price of having children P_N, and thus encourage investments in quantity. But health and schooling subventions also lower the cost of child quality, and if Z_E and Z_S, on the one hand, and Z_N, on the other, are substitutes, then the consumption of Z_N decreases.

Here, α reflects unobserved heterogeneity in the sample population. For instance, allocation of resources for public programs may not be randomly planned, and may be determined by underlying areal factors such as health endowments. A likely government placement rule is that the pattern of distribution benefits poor households (those with less endowments). Alternatively, households may sort themselves in response to availability of programs or variations in prices for preferred inputs. The association between government investments and child services provides unbiased estimates of program inputs only if the program resources are allocated independently of α or if there is control for α.

Consider a linear static demand model that accounts for such heterogeneity — unobservable individual, household and community variables correlated with the explanatory variables — in the estimation of the impact of government services:

$$\text{(8)} \quad z_{it} = \underset{1xK}{\beta'} \underset{Kx1}{g_{it}} + \underset{1xL}{\gamma'} \underset{Lx1}{r_i} + \alpha_i + \mu_{it} \qquad i = 1, \ldots, M; \quad j = 1, \ldots, T$$

where β is a K x 1 vector of coefficients of time-varying observable variables $g_{it} = (g_{1it}, \ldots, g_{Kit})$,
γ is an L x 1 vector of coefficients of time-invariant observable variables $r'_i = (r_{1i}, \ldots, r_{Li})$,
μ_{it} is the error term, $\sim N(0, \sigma_\mu^2 \mid g_{it}, r_i)$, assumed uncorrelated with g, r, α, where
α_i is a time-invariant latent variable, distributed independently across individuals (variance σ_α).

The problem under consideration is that α_i is *possibly* correlated with g and r, so that $E(\alpha_i \mid g_{it}, r_i) \neq 0$. If so, OLS and GLS estimates of β and γ will be biased and inconsistent.

Stacking (8) over i and t, the model in matrix form is

$$\text{(8a)} \quad \underset{MTx1}{Z} = \underset{MTxK}{G}\underset{Kx1}{\beta} + \underset{MTxL}{R}\underset{Lx1}{\gamma} + \underset{MTx1}{\alpha^*} + \underset{MTx1}{u},$$

where

$$Z = \begin{bmatrix} Z_1 \\ Z_2 \\ . \\ . \\ . \\ Z_N \end{bmatrix}, G = \begin{bmatrix} G_1 \\ G_2 \\ . \\ . \\ . \\ G_N \end{bmatrix}, \quad R = (\underset{MTxM}{I_M} \otimes e) \underset{MxL}{r}$$

$$\alpha^* = \underset{MTxM}{(I_M} \otimes \underset{Mx1}{e)\ \alpha}, \qquad r' = (r'_1, r'_2, \ldots, r'_M),$$

$$z'_i = (z_{i1}, z_{i2}, \ldots, z_{iT}), \quad \alpha' = (\alpha_1, \alpha_2, \ldots, \alpha_N),$$

$$\underset{TxK}{G_i} = \begin{bmatrix} g_{1i1} & g_{2i1} & \cdots & g_{Ki1} \\ g_{1i2} & g_{2i2} & \cdots & g_{Ki2} \\ \cdot & \cdot & \cdot & \cdot \\ \cdot & \cdot & \cdot & \cdot \\ \cdot & \cdot & \cdot & \cdot \\ g_{1iT} & g_{2iT} & \cdots & g_{KiT} \end{bmatrix}, \quad r'_i = (r_{1i}, r_{2i}, \ldots, r_{Li}),$$

$$e' = (1,1,\ldots,1), \qquad u' = (u_1, u_2, \ldots, u_M), \text{ and}$$

$$u'_i = (\mu_{i1}, \mu_{i2}, \ldots, \mu_{iT}).$$

$\otimes$ denotes the Kronecker product and I_M denotes the $M x M$ identity matrix. Additionally,

$$Eu_i = 0,\ Eu_i u'_j = \sigma^2_\mu I_t, \text{ and } Eu_i u'_j = 0 \text{ if } i \neq j.$$

The model in alternate form is

$$\text{(8b)} \quad \underset{MTx1}{Z} = \underset{MTxK\ Kx1}{G\beta} + \underset{MTxM\ MxL}{(I_M \otimes e)} \underset{Lx1\quad Mx1}{(r\gamma + \alpha)} + \underset{MTx1}{u}.$$

In the estimation framework used in this paper, due to Hausman and Taylor (1981), the concern is to control for heterogeneity in estimating the effect, not only of time-varying variables, but also of time-invariant variables. A prerequisite for this procedure to work is to distinguish the G and R variables that do not correlate with α. The "trick" is to partition the G variables into two groups: the first group, G_A, are uncorrelated with a, and the second group, G_B, are correlated with α. More formally, $G = [\ G_A \mid G_B\ ]$ of dimension $[\ MT \times K_1 \ \mid MT \times K_2\ ]$ The same decomposition is done for the R variables so that $R = [R_A \mid R_B]$ of dimension $[MT \times L_1 \mid MT \times L_2]$.

For fixed T:

$$\text{(9)} \quad \underset{M\to\infty}{\text{plim}} \frac{1}{M} G'_A\alpha^* = 0 \qquad \underset{M\to\infty}{\text{plim}} \frac{1}{M} G_B'\alpha^* = d_G$$

$$\operatorname*{plim}_{M \to \infty} \frac{1}{M} R'_A \alpha^* = 0 \qquad \operatorname*{plim}_{M \to \infty} \frac{1}{M} R' \alpha^* = d_R$$

where d_G $(\neq 0)$ is a K_2 vector and d_R $(\neq 0)$ is an L_2 vector.

Define two orthogonal projection operators (MT x MT idempotent covariance transformation matrices)[3]:

$$P = I_M \otimes \tfrac{1}{T} ee' \qquad Q = I_{MT} - P.$$

In the traditional technique of first-differencing, the demand equation is premultiplied by Q:

$$\begin{aligned} QZ &= QG\beta + QR\gamma + Q\alpha^* + Qu \\ QZ &= QG\beta + Q(I_M \otimes e)(r\gamma + \alpha) + Qu \\ &= QG\beta + Qu, \text{ or} \\ Z &= \tilde{G}\beta + \tilde{u}. \end{aligned}$$

OLS can be applied to obtain a covariance (within-groups) estimate of β:[4]

$$(10) \qquad \hat{\beta}_w = (G'QG)^{-1}G'QZ = (\tilde{G}'\tilde{G})^{-1}\tilde{G}'\tilde{Z}$$

If T=2, this reduces to fitting the regression on the changes: $z_{i2} - z_{i1}$ on $g_{i2} - g_{i1}$. Because $\tilde{u}$ is uncorrelated with $\tilde{G}$, $\tilde{\beta}_w$ is unbiased and and consistent, no matter if α and G, R are correlated. If G were to represent time-varying policy variables (government provision of health, education and family planning services), the above covariance method would correctly assess the impact of government programs across heterogeneous households, given that they are activated not independently of location-specific and household-level endowments.

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[3] P transforms a vector of observations into a vector of group means: $PG = (1/T)\Sigma G$; Q produces a vector of deviations from group means: $QG = \tilde{G} = G - (1/T)\Sigma G$; Q is orthogonal by construction to any time-invariant vector of observations: $QR = R - (1/T)\Sigma R = 0$.

[4] First-differencing (deviations from individual means), which eliminates α_i, is a straightforward way to derive unbiased and consistent estimates of β, using panel data. Such "within-groups" (or "fixed effects") estimation, however, sweeps out all R variables, so that in the process, γ cannot be estimated.

The Hausman-Taylor Method

To obtain consistent estimates of β and γ, the Hausman-Taylor technique is used. Basically, this IV method departs from the conventional instrumental variables procedure in that the instruments are selected from *included* variables, rather than from *excluded* variables. This is made possible by removing the time-invariant component of each regressor that is correlated with the unobserved time-invariant individual effects (α_i). Any vector orthogonal to α_i can be used as an instrument.

For both β and γ to be identified in the demand equation $Z = G\beta + R\gamma + \alpha + u$, a necessary condition is that $K_1 > L_2$, that is, there be at least as many columns of time-varying, exogenous G_A variables uncorrelated with α_i as there are time-invariant, endogenous R_B variables, for instrumental variables to be generated within the system. A necessary and sufficient condition is that the matrix

$$\left[\frac{G'}{R'}\right] P_A \; [\, G \mid R \,] \text{ be non-singular,}$$

where P_A is the orthogonal projection operator onto the column space of the matrix $A = [\, Q \mid G_A \mid R_A \,]$. On the basis of equations (8) and (9), an ordinary structural equation and two reduced-form equations can be constructed to yield 2SLS estimates:

$$\text{(11)} \qquad \begin{aligned} \Omega^{1/2} Z &= \Omega^{1/2} G\beta + \Omega^{1/2} R\gamma + \Omega^{1/2}\alpha^* + \Omega^{1/2} u \\ G_B &= G_A \pi_{G1} + R_A \pi_{G2} + Q\pi_{G3} + v_G \\ R_B &= G_A \pi_{R1} + R_A \pi_{R2} + Q\pi_{R3} + v_R \end{aligned}$$

The relevant Hausman-Taylor equation is given by

$$\text{(12)} \qquad \begin{aligned} P_A \Omega^{1/2} Z = {} & P_A \Omega^{1/2} G\beta + P_A \Omega^{1/2} R\gamma \\ & + P_A \Omega^{1/2}\alpha^* + P_A \Omega^{1/2} u \end{aligned}$$

where the MT x MT non-singular matrix $\Omega^{1/2} = \Theta P + Q = I_{TM} - (1-\Theta)P$ transforms the disturbance covariance matrix — from the model, $cov\ (a^* + u \mid G, R) = \sigma^2_\mu I_{MT} + T\sigma^2_\alpha P$ — into a scalar matrix. (Note: $\Theta = [\, \sigma_\mu{}^2 / \sigma^2_\mu + T\sigma^2{}_\alpha)]^{1/2}$). A consistent estimate of s^2_μ of the variance component σ^2_μ is derived from the within-groups regression. $\operatorname{plim}_{m \to \infty} s^2_\mu = \operatorname{plim}_{m \to \infty} \frac{1}{M(T-1)} u'Qu = \sigma^2_\mu$. A consistent estimate s^2_α of variance component σ^2_α is obtainable whenever consistent estimators for both β and γ are available.

Letting $s^2 = (1/M)(PZ - PG\hat{\beta}_w - R\hat{\gamma}_w)'(PZ - PG\hat{\beta}_w - R\hat{\gamma}_w)$, plim s^2 = plim $\frac{1}{M}(a^* + u)'(a^* + u) = \sigma^2_\alpha + \frac{1}{T}\sigma^2_\mu$. It follows that $s^2_\alpha = s^2 - \frac{1}{T} s^2_\mu$. s^2_α and s^2_μ are used in estimating Θ.[5]

Asymptotically efficient OLS estimation of β and γ (designated $\tilde{\beta}^*$ and $\hat{\gamma}^*$) in the above equation involves transforming the demand equation by $\Omega^{1/2}$, equivalent to $(1-\Theta)$ - differencing the data:[6]

$$z_{it} - (1-\Theta)\, z_{i.} = \beta'\, [g_{it} - (1-\Theta)\, g_{i.}] + \Theta \gamma r_i + [\mu_{it} - (1-\Theta)\, \mu_{i.}]. \tag{13}$$

If the time-varying exogenous variables are transformed into deviations from their corresponding time-means, the transformed variables become instruments, likewise uncorrelated with α_i. Thus, G_A have dual roles: they are used in estimating their own coefficients, and as the appropriate instruments for R_B. The time means of time-invariant variables that are not correlated with α can also be used as instruments.

Panel Data and Selection of Variables

The empirical analysis relies on restrospective information. The sample consists of randomly chosen households from the multipurpose surveys in the Bicol Region, conducted in 1978 and again in 1983. As a 2-wave panel, the Bicol Multipurpose Survey (BMS) serves as the baseline information for assessing the impact of the Bicol River Basin Development Program. The sample was drawn proportionately among

[5] See Hausman and Taylor (1981), pp. 1383-1384.

[6] Since Ω is unknown, a consistent estimator, $\hat{\Omega}^{-1/2}$, can be derived, given initial consistent IV estimates of β and γ. Estimates involving either $\hat{\Omega}$ or Ω (if it is known) have the same limiting distribution (Hausman and Taylor, 1981, p. 1387).

urban and rural populations in each of twenty "integrated development areas" in three river basin provinces. In 1978 the sample consisted of 1903 households drawn from 100 barangays; 1650 households were resurveyed in 1983 using a similar interview format. Those who migrated out of the survey area between 1978-83 were excluded from the 1983 interviews (Guilkey, *et al.*, 1988). They were replaced by substitute households.

The Choice Variables: Child Services

The variable descriptions and descriptive statistics are shown in Table 1. The household model has two sets of dependent variables: child health and nutrition (height-for-age (Z_s) and weight-for-age (Z_e)) and child schooling (average schooling attainment (Z_E)). Number of children (Z_N) or a comparable fertility measure is not included because it is not suitable for panel estimation.[7] Family planning effects may not signify current birth rate reduction outright, but better intertemporal control over patterns of birth (Rosenzweig and Wolpin, 1986).

Child Health and Nutrition

The health indicators utilized in the analysis are *z-score of height-for-age*, and *z-score of weight-for-age*, both averaged at the household level, of preschool children.[8] Preschoolers are among those most susceptible to the greatest permanent damage from protein-energy undernutrition (Paqueo, 1976; Martorell, 1982).

[7]Both quality variables considered here — education and health — and even the absent quantity variable, number of children — are by definition determined by an identical set of explanatory variables. This formulation makes it possible to ascertain whether the dependent variables move in the same or opposite direction, that is, whether they are substitutes or complements.

[8]Most children in the Bicol area start schooling at ages 8 and 9. Thus 7-year olds are considered preschoolers in this paper.

Table 1 - Definition and Descriptive Statistics, Selected Variables

Variable	Description	1978 Mean	1978 Standard deviation	1983 Mean	1983 Standard deviation
DEPENDENT VARIABLES					
Height for age	Z-score, height for age of preschool children, household mean	-2.14	1.64	-2.15	1.27
Weight for age	Z-score, weight for age of preschool children, household mean	-1.65	2.46	-1.91	0.78
Grade attainment	Age-and gender-standardized years of schooling, all children aged 6+	1.03	1.76	1.07	1.04
PROGRAM VARIABLES					
Doctors	Number of government physicians per thousand barangay population	0.49	1.20	0.33	0.63
Nurses	Number of government nurses per thousand barangay population	1.04	1.78	0.65	1.28
Midwives	Number of government midwives per thousand barangay population	0.68	0.97	0.87	1.29
Nutritionists	Number of government nutritionists per thousand barangay population	0.38	0.61	0.51	1.33
Maternity clinics	Present in barangay = 1; otherwise = 0	0.02	0.15	0.07	0.25

Table 1 - continued

Variable	Description	1978 Mean	1978 Standard deviation	1983 Mean	1983 Standard deviation
PROGRAM VARIABLES					
Day care centers	Present in barangay = 1; otherwise = 0	0.26	0.44	0.47	0.50
Rural health units	Presence (=1) of rural health unit or barangay health station in barangay	0.40	0.49	0.65	0.48
Primary schools	Number of public elementary schools with at least 4th grade level per thousand barangay population	0.85	0.68	0.61	0.50
Secondary schools	Number of public secondary schools per thousand barangay population	0.11	0.34	0.07	0.20
Family planning motivators	Number of government family planning motivators per thousand barangay population	1.08	2.15	0.79	1.47
OTHER VARIABLES					
Mother's schooling	Years of mother's schooling	5.31	3.06	5.31	3.06
Father's schooling	Years of fathers's schooling	5.39	3.12	5.39	3.12
Barangay-poblacion distance	Number of kilometers from barangay point to poblacion center	5.46	6.10	5.46	6.10

Number of observations = 669, except for height for age (307) and, weight for age (309).
Not shown: household variables - mother's age, household wealth, percentage of non-residential household members, backyard gardening; community variables - price of rice, price of milk, electricity, irrigation, urban location.

The height-for-age sample consists of a panel of 307 households, and the weight-for-age sample, 309 households. In both samples, only children for whom height and/or weight data were missing are excluded. If the missing or excluded data are correlated with child health, there would be bias in the estimates although this is assumed to be negligible. The z-score[9] for an individual child is [10]

$$\frac{\left[\begin{bmatrix}\text{height}\\ \text{weight}\end{bmatrix}\text{ of subject}\right] - \left[\text{median value of reference population}\begin{bmatrix}\text{height for age}\\ \text{weight for age}\end{bmatrix}\right]}{\left[\text{1 standard deviation from median}\begin{bmatrix}\text{height for age}\\ \text{weight for age}\end{bmatrix}\text{ of reference population}\right]}$$

The reference population used here is drawn from data assembled by the US National Center for Health Statistics and covers two child populations in the US.[11] Because international standards are known to be close to anthropometric measures for well-nourished children in LDCs, but not for poorly-fed children, the use of western growth data is justified only as approximations and not as targets (Martorell, 1982; Strauss, 1988).

Height-for-age[12] is a measure of stunting, or chronic malnutrition. It is a stable measure of long-term child health developments, since it reflects the cumulative nutritional and morbidity experience in the individual child (Ho, 1982). Weight-for-age reflects both current (acute, wasting) malnutrition and past (chronic, stunting) malnutrition (Martorell, 1982).

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[9] If a child's z-score is -2 to less than -1, he is only mildly malnourished; if it is -3 to less than -2, he is moderately malnourished; if it is less than -3, he is severely malnourished. These ranges represent first-, second- and third-degree malnutrition, respectively. This rule of thumb is recommended in Horton (1986).

[10] As recommended by the World Health Organization, See WHO (1983), p. 24.

[11] WHO, *op. cit.*, pp. 61-62

[12] Averaging the z-scores for height-for-age and weight-for-age across preschoolers in the household poses estimation problems. Strictly speaking, only within-child estimators yield consistent estimates of the effects of non-random program placements, given parental birth-spacing response to post-natal random shocks and differing sibling endowments (Rosenzweig and Schultz, 1983a; Rosenzweig and Wolpin, 1986). Child-specific estimation could not be done, however, because of data limitations.

Child Schooling

The child schooling measure, which is age- and gender-standardized, is *grade attainment*. For children i in family j, the grade completed is

$$\sum_{i=1}^{n_j} \frac{\text{years of education}_{ixsj}}{\text{mean years of education}_{ixs}} \Big/ N$$

where N is the total number of children in the household, x is the age of children and s is their gender. The index measures the *actual* attainment over the average educational attainment of children, by age and sex. This index,[13] as a convenient simplification of household demand for child education, implicitly accounts for such factors as age when started school, time in studying, days absent or temporary leaves to the degree that they increase or decrease the chances of repeating a grade; it also reflects cumulative household decisions about resource allocations on child schooling (Birdsall, 1982).

A panel of 669 households with children was used for program assessment. All children in each sample household, excluding those for whom schooling information was missing, were included in the computation of the schooling index.

The *right-hand-side* variables consist of community-level and socioeconomic background variables.[14] The data set merges household-level retrospective information on child quality and observed household characteristics with community-level policy initiatives and other explanatory variables.

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[13]As a continuous variable, the schooling index makes an implicit assumption that schooling is normally distributed, ignoring the fact that in most developing countries, education is discrete and non-normally distributed, with major nodes at grades representing elementary, high school and college completion levels. King and Lillard (1983) resolve this problem by constructing an ordered polychotomous model of schooling choice (applied to the Philippines using 1978 BMS data) with censoring.

[14]Included in the statistical implementation of the Hausman-Taylor method but excluded in this particular study are the socioeconomic background variables age of mother and household wealth, and the community variables price of rice and milk, electricity and irrigation, urban location, backyard gardening and "shadow household" (percentage of non-residential household members). Parental schooling, distance and urban location are the R variables within the Hausman-Taylor specifications.

The Program Variables

The program measures include *family planning services, health services*, and *educational services*. These governmental initiatives, most of which are provided free or at minimal cost, are utilized to proxy the price variables, since direct price variations are not available in the data set. The program variables are assumed to be correlated with latent individual and community effects. Public clinics in the Philippines, specifically, are targetted to poor, less-endowed groups, and are thus expected to serve poor households in proportions greater than their presence in the population (Akin, *et al.*, 1985).

Program variables reflect not only local exogenous constraints on health care, family planning and school availability, but also the communities' preferences for and priorities among alternative social services. Allocative responses of households, given differing health, ability and/or education, differ with respect to changes in government incentives (e.g., subsidization of social services) — which themselves are presumably distributed on the basis of exogenous endowments in the population. This implied self-selection of public services means that correct estimates of program impacts must be derived from a model in which the government programs themselves are choice variables (Schultz, 1981).

Facilities vs. Personnel

The variables selected to represent government program subsidies are health care and family planning personnel as well as facilities. The following "personnel" indicators are used: *number of family planning motivators, number of doctors, number of nurses, number of midwives, and number of nutritionists*. Also included in this study are the following "facilities" indices: *presence of rural health units and/or barangay health stations,*[15] *presence of day care centers,*[16] and *presence of maternity clinics*. As for the schooling indicators, only the following are available: *number of primary schools and number of secondary schools.*

Each of the program services is divided by the corresponding barangay population. The per capita nature of the indices is meant to

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[15]Presence of BHSes is included in the analysis in conjunction with the presence of RHUs: a village or barangay is served by either a BHS or an RHU or both.

[16]Day care centers in the Philippines function as supplemental feeding centers for preschool children. They hardly have anything to do with work-related child care programs and are not valued for their female labor-supply effects.

retain geographic variation in prices. It is hypothesized that greater accessibility of health centers/public schools — given their function of lowering the costs of child health/education or of decoding information on health or contraception — leads to higher schooling attainment, lower fertility and improved child health.

Parental Schooling

Parental education is measured as years of schooling completed. The *schooling attainment of father* reflects income effects, if the father is relatively uninvolved in child health and child production. The *schooling attainment of mother* captures both efficiency and value of time effects of women.[17] Healthwise, maternal schooling measures the mother's familiarity with child care and efficient use of appropriate health inputs. Because it is not independent of income, it is also a yardstick for the ability of the household to purchase health goods. If it reduces the average cost of schooling and health, while a substitution effect is in place, its effect must be positive. Both schooling variables may be considered endogenous because they involve allocational choice in time: parents (especially mothers) desiring fewer children are more likely to make investments in their own education.

Village-to-Town-Center Distance

Distance from barangay to poblacion is a time-invariant measure of physical distance from the village central point to the town hall. It is an indirect index of deterrence to the use of government facilities, most of which are located in poblaciones. It reflects transport and travel time costs, although in an imperfect way, since it takes no account of better transportation facilities which could lower cost.

Summary of Predicted Program Impacts

All things considered, the two following hypotheses should hold: (1) the components of child quality — health and schooling — are complements, and (2) child quality and quantity are substitutes. This should be evident in the following table, which couples the public programs to the

[17]The higher the schooling of mothers, the higher their market productivity and wages, and they may face higher levels of prices of health goods, contraceptives and child services. The more efficient they are in the use of goods, however, the lower the price levels they face. The price effect is attributed to the opportunity cost of child care, which is likely to be time-intensive on the part of the mother.

price structure in the theoretical model and predicts the compensated effects of the services on the consumption of health and schooling.

Price	Program	Z_S	Z_E
		Predicted change in dependent variable	
P_C	Family planning services	+	+
P_H	Health/nutrition services (medical care)	+	+
P_N	Health/nutrition services (prenatal care)	-	-
P_E	Educational services	+	+

The provision of family planning personnel, for example, should induce investments in child health and schooling. Health care services, provided by medical and prenatal care workers, may have opposing effects, however, because they simultaneously reduce the price of both P_N and P_H.

Estimation Strategy

The demand equation can be decomposed in the following manner:

$$(14)\quad z_{it} = \beta_A g_{Ait} + \beta'_B g_{Bit} + \gamma_A r_{Ai} + \gamma'_B r_{Bi} + \alpha_i + \mu_{it}$$

The K_1 time-varying variables g_{Ait} and the L_1 time-invariant variables r_{Ai} are exogenous and uncorrelated with α_i; the K_2 time-varying variables g_{Bit} and the L_2 time-invariant variables r_{Bi} are endogenous and correlated with α_i. For the instrumental variables procedure to work, $K_1 \geq L_2$.

The initial objective is to get consistent estimates of all parameters. First, perform within-groups estimation to obtain β_w, which are unbiased and consistent for β irrespective of the correlation between g_{it} and α'. σ^2_μ is also derived from the within-groups estimation. To get consistent estimates of γ, let $d_i = z_{i\cdot} - \beta'_w g_{i.}$ and regress $\hat{d}_i$ on r_{Ai} and $\hat{r}_{Bi}$, where $\hat{r}_{Bi}$ is derived using g_{Ai} as instruments (note that r_{Bi} is correlated with α_i).[18]

[18]If r_{Bi} are dichotomous, OLS is inappropriate for obtaining $\hat{r}_{Bi}$. Instead, a discrete-response model must be used. In this study, the probit model is used because it has a standard normal distribution. In lieu of r_i, a latent, continuous variable is used.

OLS estimates of γ_A and γ_B will be consistent. These estimates of β and γ are substituted in (14) to get estimates of the variance components. The next step is to transform (14) into (13) using a transformation formulated by Fuller and Battese (1973). Denoting (1 - Θ) by λ, the estimating equation is

$$(15) \qquad z_{it} - \lambda z_{i.} = \beta' [g_{it} - \lambda g_{i.}] + \gamma' (1-\lambda) r_i + v_{it},$$

where $v_{it} = (1-\lambda)\alpha_i + [\mu_{it} - \lambda\mu_{i.}]$. g_{it} is uncorrelated with v_{it}. Note, however, that the first term on the right-hand side is partitioned as $\beta'_A [g_{Ait} - \lambda g_{Ai.}] + \beta'_B [\hat{g}_{Bit} - \lambda g_{Bi.}]$, and the second term as $\gamma'_A (1-\lambda) r_{Ai} + \gamma_B (1-\lambda) r_{Bi}$, because the original variables g_{Bit} and r_{Bi} are correlated with α_i. $\hat{g}_{Bit}$ and $\hat{r}_{Bi}$ are obtained using two-stage least squares, with r_{Ai}, $g_{Ai.}$ as instruments. $\hat{g}_{Bit} = g_{Bit} - g_{Bi} + \hat{g}_{Bi.}$, where $\hat{g}_{Bi.} = r_{Ai} + g_{Ai}$. Likewise, $\hat{g}_{Bi} = r_{Ai} + g_{Ai}$.

The Specifications [19]

The Hausman-Taylor technique is applied to both child health and child schooling models. There are three specifications. In the first specification, *all* program variables are classified as g_A variables and assumed to be exogenous. In the second specification, there is a complete switch — all the program variables are assumed to be endogenous. Schools are reclassified as g_A in the third specification, on the assumption that they are outside even partial control of households, and thus are exogenous to the household demand framework. Also, community health endowments may have no direct systematic association with the establishment of schools (unlike the placement of health infrastructure). If such is the case, schools are distributed by the government in a random fashion. Distance from barangay to poblacion, and parental schooling, are argued to be exogenous (r_B).

Panel estimates using the Hausman-Taylor technique to control for heterogeneity across households are predicted to improve the coefficients of the program variables. Ordinary least squares estimation, with the 1978 and 1983 observations pooled, is included for comparison

[19]For variables not shown in this study, availability of electricity and irrigation infrastructure are initially in the g_A classification but are moved into the g_B category in the final Hausman-Taylor specification. The other variables have the following classifications: mother's age, g_A; wealth, g_B; "shadow household", g_B; backyard gardening, g_B; price of rice and milk, g_A; and urban location, r_A.

purposes.[20] The instrumental variables and OLS specifications are subjected to a specifications test devised by Hausman and Taylor (1980).[21]

Estimates of Program Impacts and Other Results

Standard and Within-Groups Estimates

Tables 2 and 3 give the results for child health and Table 4 for child schooling. OLS estimates, in which the explanatory variables are assumed independent of α_i, are in the first column of each table. Next to the OLS column is the within-groups estimation. The data transformation sweeps out parental schooling and barangay-to-poblacion distance, among others. As already noted, these within-groups estimates are unbiased, notwithstanding the correlation between the regressors and the latent variable. Note the dramatic changes in the parameter estimates in terms of both sign and magnitude. In general, the within-groups estimates bear little resemblance to the OLS estimates, and only few variables are relatively unaltered in both magnitude and sign. Within the programs, several variables that are otherwise insignificant in the OLS column are highly statistically significant in the within-groups estimation. A somewhat reverse pattern could also be observed in that some OLS variables lose their statistical significance in the within-groups model.

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[20]Standard errors in the OLS regressions are inconsistent because they do not treat each of the components σ_α and σ_μ as a variance in its own right. All the classical assumptions on the error term are retained in the OLS estimation — no correlation with the regressors, no autocorrelation and no heteroscedasticity.

[21]The null hypothesis tested is $E[\alpha_i \mid g_{it}, r_i] = 0$. Using the within-groups estimation as benchmark since $\hat{\beta}_w$ are consistent regardless of whether the null hypothesis is valid or not), the vector $q_1 = \beta^* - \beta_w$ is constructed, where β^* are the parameter estimators of the model being tested. Under H_0, β_w and β^* are consistent, so that if q deviates much from zero [$\operatorname{plim}_{n \to \infty} q = 0$], the hypothesis is rejected and the model is misspecified. A χ^2 - test is easily constructed from q, and is given by $\chi^2 = q'[cov(q)]^+ q = q'[cov(\beta_w) - cov(\beta^*)]^+ q$, where $(\ )^+$ denotes a Moore-Penrose, or any, generalized inverse. The degrees of freedom for this specifications test depends on the number of overidentifying restrictions $(K_1 - L_2)$.

Table 2 — Regressions on Child Health

Dependent variable: height for age

Number of observations: 307

	OLS	Within Groups	Instrumental Variables (1)	(2)	(3)
PROGRAM VARIABLES					
Health and nutrition					
Doctors	0.0405 (0.472)	0.1168 (1.563)a	0.1129 (1.446)a	0.0963 (1.504)a	0.1182 (1.611)a
Nurses	0.0322 (0.679)	0.0173 (0.393)	0.0154 (0.338)	-0.0044 (-0.121)	-0.0083 (-0.197)
Midwives	0.0717 (1.161)	-0.0030 (-0.047)	0.0174 (0.286)	-0.0159 (-0.334)	-0.0287 (-0.502)
Nutritionists	0.0875 (1.349)a	0.0079 (0.117)	0.0259 (0.385)	0.0085 (0.164)	-0.0192 (-0.315)
Maternity clinics	0.1942 (0.663)	0.3048 (1.055)	0.2918 (0.996)	0.1710 (0.764)	0.1548 (0.581)
Dare care centers	-0.0724 (-0.557)	0.1530 (1.261)	0.0724 (0.592)	0.0951 (0.957)	0.1557 (1.327)a
Rural health units	-0.0135 (-0.100)	0.1152 (0.678)	0.0717 (0.460)	-0.0067 (-0.062)	0.0283 (0.211)
Education					
Primary schools	-0.2366 (-1.885)b	-0.2660 (-1.488)a	-0.2165 (-1.297)a	0.0098 (0.098)	-0.2007 (-1.061)
Secondary schools	0.3903 (1.898)b	0.5928 (1.696)b	0.4404 (1.562)a	0.0076 (0.044)	0.5095 (1.452)a

Table 2 — (continued)

Dependent variable: height for age — Number of observations: 307

	OLS	Within Groups	Instrumental Variables (1)	(2)	(3)
Family planning					
Family planning motivators	0.0131	-0.0080	-0.0044	-0.0288	-0.0069
	(0.348)	(-0.183)	(-0.105)	(-0.730)	(-0.181)
Mother's schooling	0.0324		-0.2465	-0.3812	0.6802
	(1.035)		(-0.698)	(-0.244)	(0.407)
Father's schooling	0.0159		0.1826	0.6656	-0.7842
	(0.533)		(0.648)	(0.417)	(-0.464)
Barangay-poblacion distance	-0.0119		-0.0184	0.0098	0.0049
	(-0.842)		(-0.664)	(0.140)	(0.100)
Specification test (Chi-squared values)	16.39		12.68	4.42	3.55

a: significant at .10 level, using one-tailed test
b: significant at .05 level, c: significant at .01 level
Note: t-ratios in parentheses

Not shown: household variables — mother's age, household wealth, percentage of non-residential household members, backyard gardening; community variables — price of rice, price of milk, electricity, irrigation, urban location.

Table 3 — Regressions on Child Health

Dependent variable: weight for age — Number of observations: 309

	OLS	Within Groups	Instrumental Variables (1)	(2)	(3)
PROGRAM VARIABLES					
Health and nutrition					
Doctors	0.0576 (0.545)	0.1132 (1.207)	0.0995 (1.047)	0.0370 (0.494)	0.0394 (0.533)
Nurses	-0.0095 (-0.163)	-0.0521 (-0.941)	-0.337 (-0.599)	-0.0126 (-0.299)	-0.0081 (-0.192)
Midwives	0.3124 (4.089)c	0.3456 (4.339)c	0.3258 (4.107)c	0.1514 (2.759)c	0.1547 (2.749)c
Nutritionists	-0.0894 (-1.114)	-0.1872 (-2.205)b	-0.1747 (-2.035)b	-0.0833 (-1.401)a	-0.0894 (-1.533)a
Maternity clinics	-0.0043 (-0.012)	0.2897 (0.795)	0.3700 (1.006)	0.1367 (0.532)	0.1641 (0.644)
Day care centers	-0.0335 (-0.209)	0.1278 (0.835)	0.0866 (0.560)	0.1023 (0.884)	0.1405 (1.196)
Rural health units	-0.1247 (-0.748)	0.3551 (1.659)b	0.2394 (1.146)	0.0774 (0.628)	0.0798 (0.638)
Education					
Primary schools	-0.0024 (-0.015)	-0.4188 (-1.858)b	-0.3407 (-1.521)a	-0.516 (-0.470)	-0.2771 (-1.308)a
Secondary schools	0.0667 (0.262)	0.9990 (2.266)b	0.8233 (1.927)b	0.0033 (0.018)	0.5690 (1.298)a

Table 3 — (continued)

Dependent variable: weight for age

Number of observations: 309

	OLS	Within Groups	Instrumental Variables (1)	(2)	(3)
Family Planning					
Family planning motivators	0.0061 (0.132)	0.0834 (1.511)a	0.0669 (1.214)	0.0304 (0.845)	0.0340 (0.926)
Mother's schooling	-0.0037 (-0.095)		-0.7758 (-0.752)	-0.5686 (-0.095)	-0.8535 (-0.110)
Father's schooling	0.0114 (0.311)		-0.1063 (-0.127	-0.6301 (-0.099)	-0.3190 (-0.043)
Barangay-poblacion distance	0.0082 (0.468)		-0.0367 (-0.439)	-0.0610 (-0.227)	-0.0625 (-0.233)
—					
Specification test (Chi-squared values)	28.80		14.18	12.06	10.85

a: significant at .10 level, using one-tailed test
b: significant at .05 level; c: significant at .01 level
Note: t-ratios in parentheses

Not shown: household variables — mother's age, household wealth, percentage of non-residential household members, backyard gardening; community variables — price of rice, price of milk, electricity, irrigation, urban location.

Table 4 — Regressions on Child Schooling

Dependent variable: years of schooling — Number of observations: 669

	OLS	Within Groups	Instrumental Variables (1)	(2)	(3)
PROGRAM VARIABLES					
Health and nutrition					
Doctors	-0.0356	0.1324	0.0041	0.1314	0.1301
	(-0.677)	(2.581)c	(0.075)	(2.252)b	(2.267)b
Nurses	-0.0274	-0.895	-0.0335	-0.0840	-0.0863
	(-0.888)	(-2.878)c	(-1.070)	(-2.371)c	(-2.498)c
Midwives	0.0413	-0.1213	-0.0158	-0.1493	-0.1483
	1.049	(-2.668)c	(0.392)	(-3.050)c	(-3.126)c
Nutritionists	-0.0117	-0.1730	-0.0539	-0.1717	-0.1563
	(-0.285)	(-3.698)c	(-1.185)	(-3.283)c	(-3.093)c
Maternity clinics	0.0255	0.1712	0.0956	0.1318	0.1195
	(0.127)	(0.840)	(0.437)	(0.569)	(0.526)
Day care centers	-0.0766	-0.1114	-0.1032	-0.0915	-0.0782
	(-0.875)	(-1.259)	(-1.171)	(-0.928)	(-0.808)
Rural health units	-0.0024	0.0684	0.0345	0.1054	0.0555
	(-0.027)	(0.550)	(0.339)	(0.857)	(0.448)
Education					
Primary schools	0.1097	0.1804	0.1212	0.1281	0.1588
	(1.435)	(1.585)a	(1.178)	(1.169)	(1.242)
Secondary schools	0.1450	-0.0094	0.1999	-0.0721	0.0579
	(0.955)	(-0.037)	(1.147)	(-0.297)	(0.280)

Table 4 — (continued)

Dependent variable: years of schooling — Number of observations: 669

	OLS	Within Groups	Instrumental Variables (1)	(2)	(3)
Family planning					
Family planning motivators	-0.0068 (-0.294)	0.0048 (0.186)	0.0018 (0.066)	0.0067 (0.232)	0.0142 (0.502)
Mother's schooling	0.0357 (1.881)b		0.1650 (0.953)	0.1370 (0.350)	-0.3859 (-0.950)
Father's schooling	0.0169 (0.909)		-0.1291 (-0.723)	-0.2384 (-0.547)	0.5201 (1.015)
Barangay-poblacion distance	-0.0095 (1.128)		-0.0175 (-1.275)	-0.0295 (-0.947)	(0.0177) (0.543)
Specification test (Chi-squared values)	35.04		20.88	4.24	4.91

a: significant at .10 level, using one-tailed test
b: significant at .05 level; c: significant at .01 level
Note: t-ratios in parentheses

Not shown: household variables — mother's age, household wealth, percentage of non-residential household members, backyard gardening; community variables — price of rice, price of milk, electricity, irrigation, urban location.

Comparing the OLS and within-groups estimates as a whole, the Hausman-Taylor specifications test yields $\chi^2_{13} = 16.39$ for the height-for-age equation and 28.80 for the weight-for-age equation. The corresponding specifications test result for child schooling is 35.04. The χ^2 statistics are quite high in relative terms, signifying that the correlation between the regressors and α_i could be purged or lessened.

Instrumental Variables Estimates

The instrumental variables estimates are presented in the last three columns of each table. At a glance, the parameter estimates are quite close to the results in the within-groups estimation. There are few changes in both sign and magnitude, and most of the estimates seem robust to changes in the specification. Using the χ^2-test, with the within-groups estimates as benchmark, the height for-age equations in the last seven columns have a χ^2 statistic that ranges from 3.55 to 12.68. For the weight-for-age model, the range is from 10.85 to 14.18. These chi-squared values cut the correlation by more than half, a dramatic improvement by any standard. The χ^2 statistic for the three IV child schooling specifications is anywhere from 4.24 to 20.88. In all cases, the hypothesis of no correlation between the explanatory variables and the latent individual and community effect would not be rejected. The χ^2 statistic is highest for the first specification, as expected; the program variables are thus not exogenous.

The test statistic is lowest for the specification in which most or all program variables are classified as G_2. The model seems to perform best when most of the program variables are argued to be endogenous, thus confirming the assumption about the non-random placements of government program personnel and the close interaction between the program subsidies and characteristics of the users of health care and education services (although this should be qualified by the fact that the study has not exhausted all possible specifications). Where specifications (2) and (3) differ is in the exogeneity of the supply of schools variables, but the difference in the χ^2 statistic between them is not substantial, and it is unclear from the results whether a generalization could be made on whether the number of schools variables are endogenous.

Since the empirical results tend to show that the program variables are endogenous, only the last two Hausman-Taylor specifications are examined simultaneously for both direction and magnitude of impact of the explanatory variables.

Direct Consequences: Health Effects

An important finding from the child health models is the lack of statistical significance of the availability of health centers. RHUs, maternity clinics and day care centers have no power at all in explaining child nutritional status, either in a direct or inverse fashion. Both the presence of maternity clinics and RHUs and/or BHSes in the height-for-age estimations have the opposite sign. The presence of day care centers is slightly significant in only one height-for-age equation, and its effects could be regarded as imperceptible even if they are in the right direction.

The story is different for the supply of health practitioners. The availability of doctors does seem to be positively associated with long-run nutritional status (age-for-height). The supply of midwives and the supply of nutritionists have statistically stronger effects, this time on child weight-for-age, but the results are counterintuitive — both have the wrong sign. The number of nurses is not a significant determinant of child health in all cases.

There is a plausible explanation for this situation. If the supply of health practitioners, taken independently of the supply of health clinics, represents in a general way the outreach component of the health program in Bicol, then the combined results appear to show that it is the outreach component which is bearing the brunt of the entire program. It is the deployment of health practitioners more than the presence of mostly town-based health centers that is having a profound impact on child health. But there is an important qualification. The effectiveness of some health personnel on weight-based nutritional status is being undermined by their impact on fertility. All health workers are in varying degrees involved in prenatal care as well as in child care (including well-baby care), which have opposite effects. Nutritionists, in particular, and nurses, to a lesser extent, are probably far more effective in reducing the price of child production than in decreasing the price of child health inputs. This is the conceivable reason for the strong negative association between the supply of nutritionists and nurses, on the one hand, and child weight-for-age, on the other hand.[22] This is more than counterbalanced, however, by the availability of midwives, which is

[22]Although prenatal care and child health are assumed to be program complements in this study, they may be program substitutes in other cases. The nutritional status of infants, for example, could proxy for the nutritional status of pregnant and lactating mothers (Paqueo, 1976). Because of their high nutritional requirements, expectant and lactating mothers are the second most vulnerable group in the population.

highly sensitive to weight-based health outcomes (much more than the availability of nutritionists is sensitive to maternal care outcomes), as indicated by the quite strong positive association between supply of midwives and weight-for-age.

The reverse impact on child health of nutritionists and the lack of impact of day care centers can be explained by a concatenation of several factors: (1) the sample has a sizable proportion of small but not malnourished children, who are not responsive to the feeding program; (2) nutritionists (and likewise, day care centers) cater to the most nutritionally deprived and at-risk groups, which include *not only* preschoolers, *but also* pregnant women and lactating mothers,[23] and the results imply that like nurses, nutritionists are much more involved in subsidizing the price of children than the price of child health inputs, while day care centers may be doing both with almost equal intensity (thus, the mutually cancelling effects); and (3) the food supplementation program in the barangays may be serving those least in need of it — in a previous finding, working wives who failed to collect food at a primary health care center belonged to households in which chronically malnourished children were found (Popkin, 1975).

The positive impact of midwives appear to be inexplicable, since their primary function is prenatal care and home delivery. On closer look, however, midwives may be much more sought after for their well-baby care services. Home visits by many midwives may actually be more concentrated on child care than on maternal care. Another plausible explanation is that they may be more effective in their secondary outreach task — family planning.[24] In this dual "secondary" outreach task, it is not clear which is dominant, but certainly the positive impact of the midwives' family planning involvement is adding up to the positive effects of their well-baby care services, which explains the highly significant t-ratios for midwives. The supply of midwives, however, has negligible impact on height-for-age, and has the expected sign. Over the long haul, it appears that the prenatal chores of midwives are predominant. It is equally likely, however, that genetic endowments are the overriding factor in permanent child health and nutrition (Wolfe and Behrman, 1982), beyond the influence of program interventions.

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[23]NEDA (1982), p. 75.

[24]Midwives have a long-established presence in Bicol (and in most Philippine) communities; in a context in which family planning is a relatively new concept, midwives may be better "culturally" suited for their role as family planning counselors than are family planning motivators themselves.

Direct Consequences: Public Schools

In the case of child schooling, the t-ratios for supply of primary schools are nearly significant in most of the household-level specifications, thus confirming the positive association between primary school services and child schoooling outcomes, although it is a rather weak evidence. Both primary and secondary school variables, for practical purposes, may be considered as having little explanatory power. These results are puzzling, to say the least. One plausible explanation is that, given that each variable is a ratio of number of schools to barangay population, it is implicitly capturing what is taking place in terms of quality: a low school-to-student ratio, which implies big classroom size to accommodate more students, a large number of grades per classroom, and inadequate input resources to learning (libraries, tutorial programs, school facilities). Since price is supposed to decrease with quality, a poor school environment has the effect of raising the price of schooling, thus keeping average grade attained at a low level or slowing it down.

Yet another plausible reason is that alternatives exist for households to educate their children. One need not go to school to get education. Non-formal on-the-job training may be preferred over formal schooling, especially if the former carries badly needed short-term yield for the family, at the expense, necessarily, of higher lifetime earnings and a highly transferable form of human capital derived from a lengthy schooling investment (Birdsall, 1982). This is especially plausible for farm households, which are more predisposed to invest family resources on farm training, given that an out-of-school youth has more opportunity to do farm work. In turn, increased demand for child labor raises the opportunity cost of schooling. The impetus to move to a less costly farm training would appear to be stronger at the secondary level when schooling costs more to parents because the subsidy is less and the opportunity cost of time of children is higher. Households incur the least direct costs of sending children to public schools in the primary years, when education is universal; indeed these direct costs are somewhat offset by the schools' baby-sitting function if older children are not available for child care (Ridker, 1976).

Indirect Consequences: Health, Schooling and Family Planning Inputs

Overall, five of seven health and nutrition program coefficients of each of the child health and child schooling equations appear to behave in an identical fashion, as far as direction of impact is concerned. The exceptions are coefficients for the supply of day care centers and midwives.

On the basis of this preliminary evidence on the health care program variables, the two components of child quality in this study — health and education — may be considered *partial* complements. In turn, this suggests that to evaluate the true effectiveness of each of these programs, its cross-effects must be taken into account. If assessed in relation to its own effects only, its true value would not be correctly estimated. Conversely, given that education and health are complements with respect to health and nutrition interventions, the response of child schooling to subsidization of prenatal care services (strong components of which have been suggested by the availability of nurses and nutritionists) is opposite to the response of child health to the subsidization of medical services (doctors, maternity clinics and RHUs): the cross effects are opposite in sign. Prenatal care and medical services are thus program complements. Medical services are themselves program substitutes (each decreases the other's marginal payoff). Doctors, RHUs, and midwives may substitute for each other, as this study shows. But no substitution takes place if clinic-based services (RHUs) serve mostly poblacion and city residents while outreach services (midwives) cover rural barangays.

The family planning program is a somewhat similar matter. Availability of family planning workers has little effect on both child schooling and weight-based child health outcomes. Nevertheless, the positive response of child weight-for-age to the outreach work of family planning motivators (and likewise of midwives as family planning counselors) provides indirect evidence that weight-based health and family size are substitutes — increasing the supply of family planning services has the effect of shifting family resources to short-term child health investments. In a similar vein, the presence of family planning motivators seems to have a positive but relatively weak impact on the demand for child schooling at the household level. Lowering the costs of obtaining information on reducing birth rates may lead slightly to a reallocation of resources toward child schooling, in the manner that it does toward child health. If number of children and child quality (health and schooling) are substitutes, family planning services may substitute for health care and educational services, and may complement prenatal care services.

The primary schools' counterintuitive cross price effects on child health may be explained by the schools' inability to (1) improve child health directly, despite the presence of immunization and feeding programs in public schools[25] and (2) lower the implicit costs of obtaining information about the underlying health technology. Primary schools are less effective when health information services can be obtained inexpensively elsewhere, or when they provide information known beforehand by households. Secondary schools, as the evidence shows, are more successful in lowering the price of information that enables households to be more receptive toward child health care.

Maternal Education

In the OLS estimations, the mother's schooling attainment is highly significant and has the expected positive impact on child schooling. Father's schooling is also positively associated with child schooling, but its effect is insignificant. This is a result consistent with past education studies, which credits women as being more efficient in the household production of child schooling, and assumes men to devote less time to time-intensive child-rearing activities. The IV maternal education estimates, however, are lower than, and in one instance differing in sign from, the OLS estimate. This is not surprising, since the standard estimates are possibly overstated. Standard estimates fail to control for women's endowments such as ability, motivation and knowledge (Behrman, Deolalikar and Wolfe, 1988). Mother's education captures not only efficiency effects, but taste and genetic effects, which, if taken into account, make it more difficult to interpret the welfare implications of the impact of maternal years of schooling on child quality (Wolfe and Behrman, 1982).

Wife's schooling has no noticeable effect and has the wrong sign in the health equations. Likewise, the father's educational level is not significant, and has a mostly negative impact on health status. This runs counter to the existing empirical evidence that parents (especially mothers) with more schooling tend to be more receptive to information on child health care and proper nutrition, and to translate them into practical measures much more efficiently than those with less educa-

[25]See World Bank (1984), p. 72.

tion.[26] The unpredictability and confounding character of the parental education parameters are ascribed by Rosenzweig and Evenson (1977) to the "portmanteau" nature of education, especially when its market productivity effects[27] influence or are partly embedded in, parental wage rates.

Village-to-Town-Center Distance

The distance from barangay to poblacion is inversely associated with both weight-based child health outcomes and child schooling, but not by any significant amount. The sign is more ambiguous in the height-for-age model. It appears that transport and travel costs (as proxied by distance) are relatively unimportant. They neither deter not stimulate demand for improved child quality. Akin, *et al.* (1985) argue that in Bicol, distance is not as important as often conceived to be: households do not really travel very far to get to health facilities, and distance and transportation costs may not be linearly related for different areas in the region.

Policy Implications

The estimations have shown that health education and family planning programs, as government investments that deliberately attempt to influence the allocative role of the price system in order to alter

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[26]The result is consistent with past findings if maternal education *substitutes* for prenatal care services, i.e., mothers tend to be efficient in making practical use of prenatal health and nutrition information much more than child health information. This would induce a substitution toward more children.

[27]Since current household income is not included in this study, it is not possible to isolate the impact of education on the parents' marginal product in the market (as opposed to its effect on the marginal productivity of time in non-market chores). Thus parental education proxies as well for income, and in this sense gauges the affordability of health inputs — balanced diet, adequate medical care — for the household. While an increase in income raises the demand for healthier children, it is also hypothesized as increasing the demand for more children. And it is plausible that parents perceive an excess demand for children in Bicol precisely due to poor health endowments. Assuming that child production costs do not increase with income — this is not inconceivable, due to the availability of low-cost substitute labor from other children and domestic servants, and the compatibility of female work participation with child rearing — the wage effect of parental education may operate to relax the supply-side constraints to low health and nutritional endowments, and stimulate the demand for more children (Kelley and Da Silva, 1980). This possibility, if child health and number of children are substitutes, induces a substitution away from improved child health, causing the inverse association between child nutritional status and maternal schooling.

the household demand for both child quality and quantity, may be combined in mutually reinforcing ways (alternatively, their mutually cancelling effects, if any, can be avoided) for carrying out the dual policy goal of reduced population growth and increased human capital investments.

Clinic-Based vs. Outreach Design

The findings seem to suggest that the clinic-based health care and family planning programs in Bicol need to be deemphasized at the same time that greater emphasis must be placed on the outreach segments of these programs. This is probably what ought to happen, policy-wise, to the extent that outreach practitioners seem to be more effective in solving the supply problem in the region, where often program unavailability is a major stumbling block to improved child health and reduced family size. However, if the health centers are not utilized because of the presence of the outreach field workers themselves, then the policy implications are quite different. The substitutability among the program services, whether clinic-based or outreach, implies that they are in many ways directly competing with each other. A roving nutritionist, for example, would be in competitive terms with a day care center providing supplementary feeding; or a visiting midwife would probably be preempting a maternity clinic's on-site functions — especially if households are indifferent to distance, or to travel and time costs, which seems to be borne out by the lack of impact of the barangay-to-poblacion distance variable on child health.[28] In other words, the availability of outreach practitioners is expanding *choices* as well as availability.

Further investigation is required to ascertain how much more outreach workers are just another option for securing health care, but it is at least clear that the lack of explanatory power of clinic-based services is no argument for immediately scaling them down, in the same manner that the strong impact of outreach services is no argument for hastily

[28]It is possible, of course, that unobserved community health infrastructure is being proxied by the distance variable, which would indicate that distance itself is in the same category as the other program variables. But Akin, *et. al,* (1985) provide compelling evidence that distance, transport and travel time costs, and waiting time costs have no impact on the demand for health care in Bicol.

drawing up plans for the rapid deployment of field personnel.[29] The government possibly oversubsidizes households by shouldering even costs that they are willing to pay, such as transportation costs, given their indifference to distance.

A policy area that needs rethinking is the marginal productivity tradeoff between program substitutes. The health programs can be made more effective if various outreach workers can fan out to separate populations beyond the distance "threshold" of their own health clinic bases. In family planning, for example, this would correct a situation in which outreach workers tend to confine their services to urban, electrified areas in Bicol, and thus miss out on women in poor, far-flung areas (Herrin, 1984). Otherwise, all primary health care services that theoretically can be provided in clinics must be "centralized" in health centers that are within reach of the community. Barangay health stations are widely accessible in Bicol[30] and provide the least expensive source of primary health care. BHSes are the substructures where much of the innovation should take place. The resulting mix of clinic-based and outreach health services should be much more effective in conserving medical resources, avoiding competition among the services, and encouraging a more systematic way of delivering care to the community.

Prenatal Care and Medical Care

The weak effects of many of the health care and family planning services on child health, the results suggest, are in part explained by the confounding effects of prenatal care, which somewhat neutralizes the relative gains in child quality. The obvious way out of this policy dilemma is to create a situation in which prenatal care depresses the price of children by much less than medical care lowers the price of child quality (or, the cross-price effects of health care must be made greater than the cross-price effects of prenatal care, to induce a decline in

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[29]The constant attempt of the Philippine government to fill the outreach "gap" through massive infusions of barangay health workers into all 40,000 barangays may be counterproductive, as Akin, *et al.* (1985) argues, because of its high opportunity costs: the money could be spent alternatively to upgrade the existing system of public clinics and make it more financially attractive to career-oriented medical professionals, of which the country has an apparent ample supply.

[30]BHSes, as the prime delivery points for health services, may be considered an intermediate stage between town center-based clinical services (RHUs), and itinerant professionals. As an outreach health unit, the BHS registered the highest proportion of households residing at the "most ideal" distance and the lowest proportion of households located at the "least desirable" distance. See NEDA, 1982, p. 85.

demand for more children by an amount that is greater than the decline in demand for child quality). But this is not a simple case of reallocating program resources away from prenatal care and toward medical care. In fact, a strong case can be argued for better prenatal care, since it affects child quality as well: malnourished mothers give birth to malnourished, and possibly, genetically weak, children; healthy mothers have healthy infants. The tradeoff faced by parents is between having an additional healthy child and investing health resources on each child. In practical terms, subsidizing health care programs should not compromise the quality of prenatal care. It is here that family planning holds the balance between prenatal care and health care. Family planning subsidies cause a rise in the price of children, and in a sense increases the cost of prenatal care relative to medical care. If the health care program were to be reconfigured, it should reinforce the coordination between family planning and health services, instead of pursuing both in near isolation from each other. "Piggybacking" family planning services to health services, for example, may yield more useful results in the sense that health centers can directly boost child survival and nutritional status (a precondition to the decline in demand for children).

Alternative Schooling Investments

In spite of full government subsidies, the opportunity cost of schooling may be positive due to poor quality and the existence of alternative educational investments, and to optimizing households must be higher than that of other options. The lack of explanatory power of schools on grade attainment reflects in part this low quality and in part competing schooling choices. In the latter case, parents may be responding to economic signals that favor short-term, high-yielding alternative education, like farm training, that improves the average labor productivity of the household. Indirect evidence supports this inference.

Parents may resist shifting their labor-oriented investments back into formal public education unless the social premium on education is raised.[31] To encourage parents to set their sights on higher lifetime yields, it may be necessary to undertake employment-generating proj-

[31]The government can marginally influence the family's allocational decisions by prescribing premature entry of children in the labor force, or passing compulsory school attendance laws, but these could be set aside as counterproductive; too often they penalize the poor households which can barely survive without the income contribution of their children.

ects that require the hiring of persons with a prescribed minimum level of skills. The accent on skilled labor is important, since efforts to increase agricultural or industrial productivity might raise the demand for skilled and unskilled labor simultaneously, thus keeping the demand for unskilled farm child labor at a high level, and the demand for more schooling at a low level. At any rate, these are policy prescriptions that entail long gestation periods, and policymakers must deal with the high information costs to parents of evaluating the advantages and uncertainties of various alternatives.

The Efficiency and Wage Effects of Parental Education

The multivariate analyses indicate that policies aimed at increasing child schooling will not benefit much from an improvement in the educational levels of parents, mothers particularly. There seems to be limited scope for intervention, in this case, although this is not unexpected if, indeed, the impact of maternal schooling on child education has been substantially overestimated in past studies. To enhance child schooling, policymakers must look elsewhere for guidance, and the policy measures already suggested to discourage parents from making low-yielding investments in farm training are examples of what can be done to solve this problem.

Neither is there confirmation in the regressions of a strong relationship between parental schooling and child health. In this case, holding adult education classes on nutrition for mothers might bring about the payoff that is missing in formal maternal schooling investments.[32] Such training might induce substantial and immediate improvements in the mothers' ability to interpret health information

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[32]Given the confounding effects of parental schooling, however, attempts to deal with parental education independently of other policy measures are not likely to be very successful, and that it might be more worthwhile to look for so-called "proximate" variables that mediate between education and child quality. For instance, if the negative association, no matter how small, between mother's schooling and child health is explained by the pronatal impact of more years of mother's schooling, then the compatibility of female work force participation with child rearing ought to be looked into. An increase in the wage effect of maternal education would not necessarily increase the price of time of mothers, due to the non-competitiveness of available female jobs with child care functions, the presence of inexpensive surrogates, or even an increase in non-earned income. The key, in this case, is to introduce policy interventions, in conjunction with investments on parental education, that effectively raise the cost of child rearing. This would require concurrent complementary moves, such as enhancing women's opportunities in the market (creating jobs that are incompatible with child-rearing) and promoting development measures that reduce the need for unskilled child labor (such as agricultural mechanization) or induce parents to invest in child schooling (such as employment-generating industrial projects).

and to be more efficient in the production of health inputs — changes that enhance household incentives to spend more resources in their existing children (in the process increasing the perceived costs of children). An increase in mother's efficiency also improves preventive care, causing the indirect costs of medical services to fall.

Summary and Conclusion

This paper has investigated the direct and indirect consequences of public services in health, education and family planning on child health, schooling and indirectly, fertility, using a reduced-form household demand model that (1) merges household-level and community-level information from the Philippines, and that (2) relies on an instrumental variables estimation technique which resolves the combined problem of unobserved effects and the endogeneity of the program (and other) variables. The evidence from the IV results, considered more reliable than standard regression results, suggests that households do respond to variations in the price of governmental programs by shifting the allocation of family resources from an assured number of children to less but healthier and better educated children.

If the results can be generalized, they support the hypothesis that a combinatorial program (that considers both own- and cross-effects) is more effective in achieving the desired policy of reducing population growth and increasing human capital.

The recourse to the Hausman-Taylor technique, using the 1978 and 1983 Bicol surveys, shows that had the standard estimates been used to make inferences on substitution and income effects, a different set of conclusions and policy consequences would have been arrived at. For example, the use of the OLS coefficients of many of the supply of health practitioners variables would have led to the premature rejection of the alternative hypothesis that these regressors are strongly associated with either health or schooling, and the unavoidable conclusion would have been that the entire primary health care system in Bicol is by and large ineffective, regardless of type of facility or personnel. That would have run counter to the evidence presented in this study.

All things considered, the findings should stimulate more cost-effective means of implementing policy measures in health, education and family planning. The often knee-jerk response of the government to fill "gaps" in eduction and health care delivery — pour more resources that increase the supply of educational facilities, or health field personnel — may be a costly way to address the problem of stimulating demand in these areas. Given the lean and overstretched capacity of the government to provide resources, it makes more economic sense to redeploy existing resources in a manner that maximizes their impact, direct and indirect, on health, education and fertility. A carefully designed policy that is more sensitive to child health and schooling responses to variations in the prices of government programs — including those not directly related to health, schooling and fertility reduction — will be much more effective in improving household welfare in the context of the national goals of reduced population growth and improved productive stocks of human beings.

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