Diminishing Returns to Scale in Family Planning Expenditures:

Thailand, 1976-1981*

T. Paul Schultz

Yale University

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Abstract

This paper assesses the effectiveness of government subsidies to public and private family planning delivery systems to reduce fertility in Thailand before 1980. Fertility rates declined by half in Thailand from 1960 to 1980, and individual based estimates from regression and poisson models for birth rates during 1975-1980 confirm that a significant share of the variation in fertility is associated with the provincial level of expenditure on family planning. Additional hypotheses are tested for which segments of the population (i.e. by age, education, rural/urban, farmers, etc.) benefited most from local family planning programs activities, whether the private and public programs are substitutes or complements in each other in the production of more effective birth control, and the differential returns to scale in the private and public delivery of family planning. Although limitations of the data and the analysis are noted, the findings strongly suggest that the rapid increase in female education and the range of public subsidies extended to voluntary family planning programs in Thailand were both important in accounting for the rapid decline in that country's fertility.

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I. Introduction

Between 1960 and 1980 total fertility rates in Thailand declined by nearly half. This sharp and continuing drop in fertility may be attributed to developments associated with the country's rapid economic growth (6-8 percent per year) and related to a variety of family planning programs that subsidized both public sector provision of services and "private" delivery of birth control information and supplies through nonprofit organizations. There are few estimates of the effects of family planning program expenditures on the decline in fertility for any country (Stolnitz, 1987), and Thailand is no exception. Public services in most parts of the world are being retrenched and refocussed on only those activities that the state performs more efficiently or more equitably than the private sector (World Bank, 1988).

This paper assesses the effectiveness of public subsidies to public and private delivery systems and tests a variety of hypotheses as to how the effect of public expenditures on family planning programs differs across socioeconomic groups and whether the marginal effect of public expenditures changes with the scale of programs. Without a consensus on the appropriate methods to use to estimate the cost effectiveness of family planning, the current round of reductions in public sector programs is likely to reduce priorities for family planning at a time when population growth continues in Africa at nearly 3 percent per year and is still 2 percent per year in ports of South and West Asia.

The analysis relies primarily on the Thai 1980/81 SocioEconomic Household Survey (SES) which is merged with 1975 data on province-level expenditures on family planning.

Recent fertility rates are estimated at the individual level by "own-child methods" and fit to covariates by linear instrumental variable techniques and by maximum likelihood methods based on a time homogeneous Poisson process (Hausman et al., 1984; Maddala, 1983; Mullahy, 1986). The estimates permit us to inquire: (1) How much of the fertility decline by age groups can private and public programs explain? (2) Are the returns to scale in the two programs constant? (3) Are the private and public sector programs complements for each other or substitutes? (4) For which segments of the population (i.e., educated, urban, farmers, etc.) is each program more influential?

Controls for the influence of economic constraints on fertility, such as the mother's education and household's permanent income level (per adult), are also included in the model. The SES was designed to obtain accurate data on consumer market expenditures, household consumption of own production, and income in kind. Because consumption is a better proxy for permanent expected lifetime income than current reported income, a comprehensive measure of total household consumption per adult is used to approximate the household's expected income constraint. The effect of household income on fertility decisions is estimated by instrumental variable methods because realized household income is endogenous to the fertility determination process, most notably because it includes the effect of the woman's working decisions in the labor market. Municipal and regional divisions of the Thai population as well as the economic and educational characteristics of individuals are interacted with local planning program expenditures to assess how these public and private program expenditures affect distinct segments of the population.

II. An Empirical Model: Explanatory Variables and the Data

Four issues in research to evaluate population policies are discussed in this section. First, the nature of the policy intervention is described and the modeling issues for empirical study are narrowed. Second, the environmental constraints on reproductive goals and achievements, other than the policy intervention itself, are characterized in order to control for them statistically. Third, a measure of the policy objective is proposed that is a proxy for the period fertility rate. Fourth, a statistical model is presented to approximate the association between a non-negative integer measure of recent fertility and the policy intervention and controls.

The Family Planning Program

Family planning programs are designed to help couples control their fertility. A simple measure of the effectiveness of public expenditures on family planning is their impact on current fertility.[1] A program might affect fertility by a variety of routes: (i) subsidizing clinic-based services of physicians and nurses and contraceptive supplies required by contraceptors or sterilizers; (ii) subsidizing the distribution of simply birth control supplies and information provided by paramedics such as midwives, who canvas local communities to encourage adoption and use (i.e., outreach); and (iii) disseminating information about birth control options and where to go to obtain specific methods. Whether these program functions are best administered within the public sector, or should be privatized through subsidies and incentives to existing nonprofit organizations and private health care providers, is a subject of current debate (World Bank, 1988). The role of the public sector in family planning may also appropriately change as the demographic transition progresses. For example, initially, the program must diffuse basic knowledge about a new and innovative technology of birth control, whereas at later stages the program provides a routine service and source of subsidized contraceptives to an informed and receptive public. More successful family—planning programs in low-income countries have generally included all three activities—clinics, outreach, information but it is typical for them to be integrated to some degree. Consequently, evaluation of the separate segments of the program is often difficult, if not impossible.

Such public sector family planning programs might exhibit several properties. First, they are eventually subject to diminishing returns to scale as they expand, where scale is defined as the level of local programs expenditure per potential beneficiary, or in this case public outlays per woman of childbearing age (Schultz, 1988). Second, separate program activities such as clinic subsidies and outreach are likely to be substitutes for each other. In other words, each reduces the other's marginal payoff. This potential for substitution can be mitigated to the extent that the clinic and outreach segments of the program actually serve separate populations, such as when the urban population is served by the hospital and clinic facilities, and the rural population is served by the paramedic outreach activities. In this case, clinics and outreach may not erode the marginal product to the other program segment. Third, information may either complement or substitute for the effectiveness of the other two distribution networks (Schultz, 1988). Fourth, to the extent that education lowers the cost for couples to evaluate new birth control technologies, the informational aspect of the program is less valuable to more educated couples and hence less effective in adding to their birth control capacities (Rosenzweig and Schultz, 1988). Finally, the greater the household's permanent income per adult, the less influential may be the financial subsidies provided by the program for adoption and supplies.

Other Factors Affecting Fertility: Control Variables

Household demand theories of fertility assign importance to the opportunity cost of the mother's time employed in child care. In Thailand, where women participate frequently in the market labor force and are also primarily responsible for child care (Knodel et al., 1988), it is expected that when they can earn a higher wage, their income and child costs increase. It is generally observed that the net effect of these offsetting income and substitution effects associated with an increase in women's wage rates is to reduce the desired and achieved levels of fertility (Schultz, 1973, 1997). Among the four thousand women in Thailand who reported a wage or salary in the 1981 SES, an additional year of schooling is associated, on average, with a 26 percent increase in their hourly wage, similar to the proportional association between education and wage rates among men.[2] Women's years of education in Thailand is expected to decrease fertility, but this effect may vary by levels of schooling if years of education affected fertility only to the extent that it increased market wage opportunities and the value of women's time. The household demand framework applied to the analysis of fertility suggests that the wage rate available to men should exert a less negative (and perhaps even a positive) effect on fertility compared with the effect of wage rate of women (Schultz, 1973, 1997). Nonearned income should increase the desired number of children, for increases in physical wealth that add to nonearned income relax the household's budget constraint by raising the opportunity cost of the time of either parent in child care without necessarily increasing the price of children.

It is possible to disaggregate income sources for most households from the 1981 Thai SES, particularly those in which both spouses work together in self-employment, such as agricultural households which constitute the largest group of families in Thailand. Thus, measured household income distinguishes the earnings of all adults in the nuclear family[3] and nonearned income from land and other assets. From most other studies of fertility determinants in primarily agricultural populations at Thailand's income level, family income from land, other assets, and husband's earnings are generally associated with higher fertility (Schultz, 1973; Mueller and Short, 1983).

Farmers incomes in some regions of Thailand are volatile, depending critically on unpredictable rain-fed crops. Transitory variations in household income are, therefore, a relatively large share of total income variation in Thai agriculture. This may also be true of other types of self-employed activity in Thailand. In approaching fertility as a household demand commodity I expect to find a responsiveness of fertility to variation in household permanent or lifetime expected income, but I do not have either the panel data or the guidance from theory on how fertility is likely to respond to transitory income variation. Using the 1981 SES, Paxson (1992) found that annual deviations in weather from historical averages explain savings (income-consumption) behavior among Thai rice farmers. Because of the relative importance of transitory income among poor self-employed, the household's total expenditures are specified as my proxy for household permanent income.

Many questions in this survey were directed to quantifying current, periodic, and durable expenditures, the imputed valued of household-produced consumption, income in kind, and the rental value of owner-occupied housing. This exhaustive enumeration of expenditures and auto-consumption is likely to be more complete in measuring consumption than are the handful of questions devoted to reporting wage and nonwage income of family members. This proxy of household permanent income is a combination of the labor productivity of men and women in the household as well as asset income. This income is divided by the number of working age adults (15-65) but not by children to avoid contamination with the fertility decision.[4] If the mother's education is also controlled in the fertility relationship, this household income variable should capture primarily the effect of male earnings and nonearned income that are more likely to be pronatal sources of family income, while the mother's education proxies for the woman's wage opportunities.

Nonetheless, this consumption-based measure of permanent household income is not exogenous to the fertility decision. As noted above, it contains the woman's market earnings and thus embodies her market labor supply decision that is determined jointly with fertility over the life cycle. This probably imparts a negative bias to the directly estimated (OLS) effect of household income (or consumption) on fertility. Second, the arrival of children, both planned and unexpected, may lead to temporary dissavings by parents to finance the consumption needs associated with young children. This source of endogeneity of savings could contribute to a positive bias in the estimate of the income effect based on household consumption. Finally, children may contribute to family income before they reach age 15, but they are included in the denominator of the family consumption per adult variable only after they are 15. This feedback effect of fertility should be negligible in the current context because the analysis deals primarily with births in the last five years, and this measure of recent fertility is not correlated with the number of own children 10 to 14 in the mother's household.[5] To avoid these several sources of simultaneous equation bias and bias due to random errors in measurement of permanent income, the household consumption effect is estimated by instrumental variables. Family owned land, unearned income, husband's age and education are all used as identifying instrumental variables to estimate the anticipated positive effect of husband's permanent earnings and family assets on fertility.

Economic theory provides little guidance on how the husband's permanent income is functionally related to fertility. With development, fertility declines and consumption increases, suggesting that any positive fertility effect of rising incomes is eventually more than counterbalanced by increases in the opportunity costs of children, although increase in women's wages may be responsible for some of this decline in fertility. A nonlinear effect of income is, therefore, not entirely unexpected, and indeed some studies have emphasized this finding (Encarnacion, 1974). Others have observed that the positive income-fertility relationship expected from demand theory is evident only in rural areas of low-income countries (Lee and Bulatao, 1983). These empirical patterns are frequently assessed, however, without controlling for women's wages or education levels. It is not possible to distinguish, therefore, whether the increasing education and wage levels of women relative to men that often occur during the process of economic development are responsible for the nonlinear income association with fertility, or, alternatively, whether the effect on fertility of male wages and nonearned income is initially positive and changes to negative as income levels increase. These possibilities are explored below.

A quadratic in age is included in the fertility equation to capture the biological capacity and behavioral tendency to have births in certain periods of the life cycle. But because more educated women delay the onset of childbearing, in part to complete their schooling, an interaction between age and education is also allowed in the subsequent empirical analysis.

Public family planning expenditures are reported by province in Thiencay Kiranandana et al. (1984: Appendix). Separate figures are given for the National Family Planning Program under the Ministry of Health and for the public support of the private Family Planning Associations (primarily Planned Parenthood Association of Thailand in 1975, though support had been extended to several additional private nonprofit groups by 1980) (Rosenfield et al. 1982)). Expenditures are reported, however, for only a few years, including 1975 and 1980. Since expenditures can only influence fertility with a lag, the 1975 level of expenditures is expected to impact on 1976 and 1981 fertility more strongly than would 1980 family planning expenditures. Empirically, the inclusion of the 1980 family planning program expenditure values, in addition to those reported for 1975, did not statistically improve the fit of the fertility model, and they are collinear (e.g., r = .5 to .70). Consequently, only the 1975 levels of program expenditure per woman 15 to 49 are reported below, although they should be interpreted as the average of the past several years of family planning expenditures in these programs. If they were available, family planning inputs for all years from 1970 to 1980 should be analyzed, employing a smoothed distributed lag specification (Schultz, 1988), but only 1975 and 1980 are currently available. Other local health and education programs might also affect the demand for children and be correlated with family planning activity levels.

Omission of these interacting health and education programs is likely to bias upward the reported estimates of how large an effect (negative) the family planning program has independently exerted on Thai fertility. Regional expenditure accounts on health and education in Thailand were not available to pursue this possibility here (Rosenzweig and Wolpin, 1982).

Finally, a dummy variable is included to indicate whether the woman resides in a municipal area. Prices and wage opportunities differ between rural and municipal areas of Thailand, and they are likely to raise the cost and lower the immediate productive benefits of a large family to the municipal resident. The municipal variable is designed to capture the effect of these unobserved variables on fertility. Approximately, one-third of the SES sample of 12,620 women age 15 to 49 live in urban or municipal areas, whereas nearly one-half live in rural villages. The remaining sixth reside in "suburban" areas called sanitary districts, which are often very similar to the urban municipal areas, but have not been so reclassified (Goldstein and Goldstein, 1978). The fertility model is subsequently estimated within these broad strata and 297 local sampling clusters to assess specification errors in the model related to omitted regional variables that influence fertility and are correlated with the core household variables: husband income and women's education.

Because different birth cohorts have experienced very different initial living conditions when they started their reproductive careers, one might expect their recent fertility to respond differently to the current provision of birth control services and to their environmental and economic constraints. There are, regrettably, no retrospective data on the life cycle conditions applicable to the couples at a common stage in their life cycle, nor sufficient published data on family planning program activity over time by region to describe how long individuals were exposed to government-subsidized modern birth control information and services.

In sum, the reduced-form demand equation for the recent period rate of fertility of a woman is assumed to be a function of local family planning service subsidies, the woman's economic constraints on her demand for children, and the biological constraints (e.g., age) on her supply of children:

$$F = F\ (X_{1p},\ X_{2p};\ A_w,\ E_w,\ I_w,\ M_w)\ ,$$

where F is a measure of a woman's recent fertility in 1981 (described below); X_{1p} and X_{2p} represent public expenditure in 1975 on family planning to The Health Ministry and on private nonprofit family planning organizations, respectively, at the local province level; A_w is the woman's age; E_w is her completed schooling; I_w is her household's consumption per adult (proxy for permanent income); and M_w is a dummy variable equal to one if she resides in a municipal area.

The previous discussion leads to the hypotheses that this fertility function will exhibit certain sign patterns in its first and second order derivatives:

(1) $dF/dX_1, dF/dX_2 < 0\ ;\ d^2F/dX_1^2,\ d^2F/dX_2^2 > 0$

and probably, $d^2F/dX_1dX_2 > 0$

(2) $dF/dA > 0\ ;\ d^2F/dA^2 < 0$

(3) $dF/dE < 0$

(4) $d^2F/dEdA > 0$

(5) $dF/dI > 0$

(6) $dF/dM < 0$

As suggested in Figure 1, estimates of the fertility function provides information on the marginal effect of the family planning inputs, dF/dX, evaluated at the sample mean of the program inputs $\bar{X}$, and the other explanatory variables, summarized in the figure by $\bar{Z}$. If family planning effort is subject to diminish returns to scale as hypothesized, the marginal effect of public expenditure at the sample mean will be less than the average effect, i.e., $dF/dX = (F_1 - \bar{F})/\bar{X} < (F_\circ - \bar{F})/\bar{X}$, and the implied total effect of the program will exceed the effect estimated by multiplying the marginal product times the average level of program expenditures.

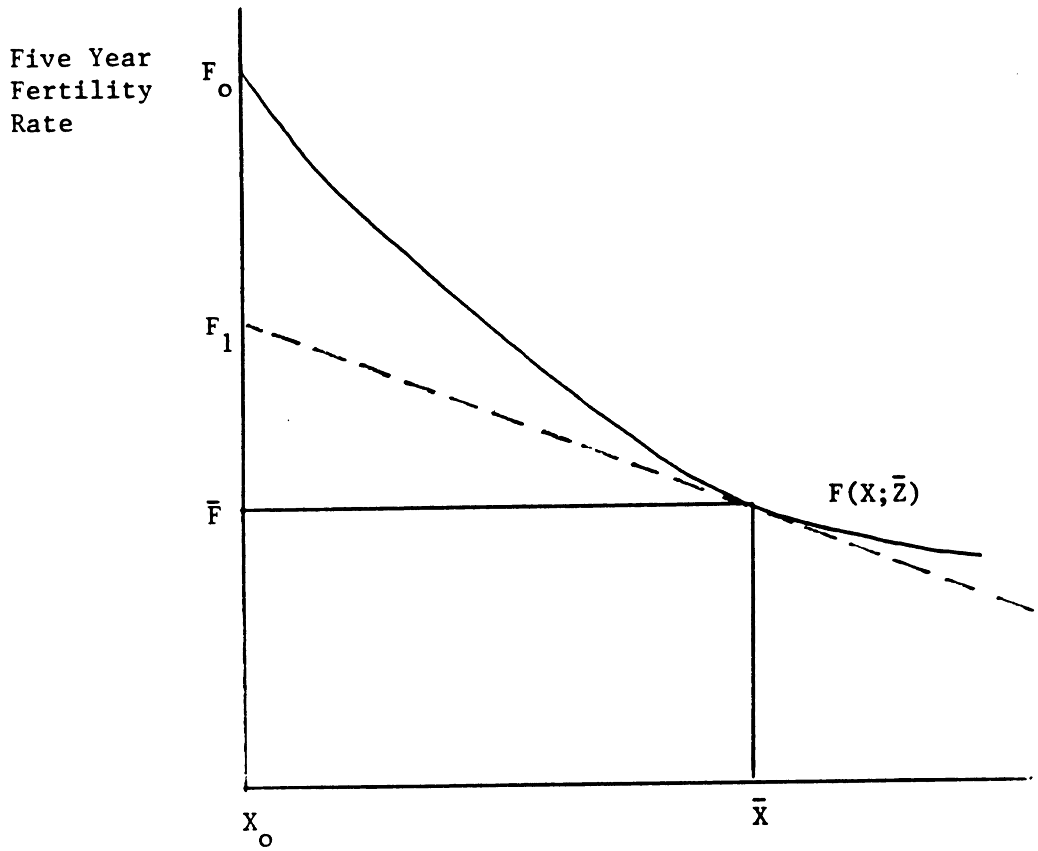
III. Measurement of Recent Fertility Levels

A limitation of many studies that seek to estimate the effect of family planning policies in low-income countries is that individual data on fertility are rarely observed in conjunction with household economic and community program variables. The World Fertility Survey, for example, collected, for many countries, comprehensive reproductive histories, but did not collect even rudimentary household income, wages or expenditure information. Further, it did not collect community-level data on policy interventions or public expenditures on relevant welfare programs that might have influenced fertility.

To obtain these three types of data together in Thailand, priority was assigned to obtaining good household expenditure and income data. However, it was not possible to disaggregate family planning expenditures below the provincial level, and an indirect measure of recent fertility, derived from the registry of household members identified by their relationship to the head of household, had to be relied on.

FIGURE 1

HYPOTHETICAL RELATIONSHIP BETWEEN LAGGED EXPENDITURES ON FAMILY PLANNING AND RECENT FERTILITY RATE



Public Expenditures on Family Planning per Woman

When vital registration systems are incomplete, as they often are in low-income countries, age-specific birth rates can be indirectly estimated by calculating the frequency of surviving own-children per woman of childbearing age in a census or representative household survey (Cho et al., 1986). This method of inferring aggregate fertility is then refined, such as by incorporating child mortality adjustments, to estimate recent period fertility rates for a country and for regions in a large country or sample. This paper examines, at the level of the individual woman, the number of her own coresidential children under the age of 5, or alternatively her own coresidential children under the age of 10, as a proxy for her recent five-year and ten-year fertility rates, respectively. This measure of fertility at the individual is not generally analyzed because fertility histories are available, but fertility surveys do not devote questions to the core economic variables required here. Because any family with its own sources of income is separately enumerated in the 1981 SES, the relationship of household members to the head is sufficient in virtually all cases to match women to their offspring. Nonetheless, two to three percent of the children under 5 years of age cannot be attributed to a mother, either because household structure information is inadequate for matching or because the mother does not live with her child (see the Statistical Appendix). A few cases also occur of polygamous unions in the same "nuclear family" that prevent matching a child to her mother. It is not clear how the unmatched children and women who are ultimately excluded from this analysis might distort the working sample and bias estimates derived from it. Their number is not large, however.[6]

One independent check on the adequacy of the child-woman fertility variable is to aggregate it up to the level of a national estimate of the total fertility rate.[7] This is done by calculating the average (surviving) fertility (children under 5) rate for women in each five-year age group, 15 to 19, 20 to 24, ..., 45 to 49. The sum of these rates, multiplied by five, is a proxy for the total (surviving) fertility rate (TSFR) during the period 1976-1981. Weighted by the four sampling strata of the survey (i.e., Bangkok, municipal areas, sanitation districts, and rural villages), this procedure yields an estimate of 2.89 surviving children per woman over the reproductive ages, if the implied age-specific birth rates remained unchanged over their reproductive lifetime. Age-specific mortality in the 1975-1985 period implies that about 4.7 percent of the children born between 1976 to 1981 would have died before the 1981 SES. A total fertility rate (TFR) of 3.03 is thus estimated. National estimates of TFR from other sources are 3.67 in 1975, 3.01 in 1980 and 2.ll in 1986 (Ratanarat, 1986; Chayovan et al., 1988).

Thus, the level of the 1981 SES-based measure of fertility in 1976-81 is not out of line with national estimates. Nevertheless, a more relevant comparison is possible at the provincial level. An estimate of TSFR can be calculated from the 1981 SES for 62 of Thailand's 72 provinces (i.e., Changwats). When this estimate is regressed by ordinary least squares on an independent province level estimate of TFR in 1980 (Kiranandana et al., 1984), the estimated factor of proportionality is .9:

$$TSFR_{P76-81} = \underset{(25.8)}{.904} TFR_{P80} \qquad , \ R^2 = .92$$

where p = 1, ..., 62 indexes the provinces, and the t ratio is reported in parentheses. The level of child mortality of about .05 could account for the slope coefficient being less than one. In sum, provincial variation in child mortality and mother-child separation rates are apparently moderate enough to encourage an analysis of TSFR as an approximation of the rate of childbearing in the Thai population from 1976 to 1981.

IV. A Statistical Specification

The number of children a woman can have under a specific age is a nonnegative integer. For most Thai women, the number of their children under five is zero. Specifically, for all Thai women 15 to 49 the mean of this measure of fertility is .38, whereas for women 20 to 29 it is .60. A model is specified linking this count of own children to a series of covariates that are assumed exogenous, or estimated by instrumental variables if endogenous, and thought to influence this reproductive outcome. Because the dependent variable measuring fertility takes this categorical form, ordinary least square (OLS) estimates of the relationship may not be efficient or unbiased. A Tobit model would accommodate explicitly the truncated (below zero) form of the dependent variable but would not incorporate the discrete positive integer distribution of outcomes. The Poisson model is an alternative framework frequently used to describe events that occur randomly and independently in time (Maddala, 1983). It is a plausible characterization for the discreteness of counts of events observed over a fixed period of time, e.g., the last five years (e.g., Ainsworth, 1988).

The likelihood that the i^{th} woman will have c_{it} own-children under age five living in her household in time period t is assumed to be described by a Poisson process:

$$f\,(c_{it}) \;=\; (e^{-\lambda_{it}}\;.\;\lambda_{it}^{c_{it}})/c_{it}\;! \qquad (1)$$

The influence of a vector of exogenous individual resources and traits, as well as local public and private family planning program services, x_{it}, on the rate of own children is assumed to take the form:

$$\lambda_{it} = e^{x_{it}\beta}, \tag{2}$$

with β a vector of parameters (Maddala, 1983). The Poisson specification implies the restriction that the mean (*E*) and variance (*V*) of c_{it} are equal:

$$E(c_{it}) = V(c_{it}) = \lambda_{it} = e^{x_{it}\beta} . \tag{3}$$

The maximum likelihood method is used to obtain estimates of β.[8] Therefore, the corrected robust standard errors are reported below.

There are three limitations to the Poisson analysis of fertility, of which the first two apply also to other approaches as well. Births do not occur independently over time; a conception-gestation period of about nine months elapses between births. The Poisson model neglects this feature of reproductive biology, as do, implicitly, linear (OLS) regressions or some other dynamic representation (i.e., Vector Auto Regression) of reproduction (Hotz and Miller, 1988). This neglect of the gestation requirement of the birth process should be a less serious limitation for the study of a low-fertility population, which Thailand has become in the last two decades, than it would be for an analysis of a high-fertility population.

The second problem is measurement error associated with the dependent variable. Own children who were born in the last five years but died or reside in another household are not counted in the analysis. The own-child measure would seem an adequate proxy for recent period fertility in populations such as Thailand where infant and child mortality rates are low and virtually all children live with their mothers.

The third limitation, specifically of the Poisson model, is the statistical restriction that the expected value of the child count is equal to the variance of the process across women.

Some applications to the frequency of accidents or medical conditions have noted that the empirical variance may exceed the expected value. Although this "over dispersion" problem can be reduced by conditioning on covariates, a corrective strategy is appropriate because over dispersion can lead to inconsistency in the standard errors of the maximum likelihood estimates of the Poisson model. In the case at hand, the count of Thai children under five in 1981 and the magnitudes of the sample variance are similar to the means. For women age 15 to 49, for example, $E(c_i) = .376$; $v(c_i) = .417$ (for other age groupings see Table A-2), whereas the residual (unexplained) variance in the model is six percent smaller than the mean, or .354. One interpretation might be that this restriction of the Poisson model is consistent with the data and "over dispersion" is not a serious problem in this study. Nonetheless, a consistent estimation approach is adopted that is robust to this form of over dispersion.

V. Empirical Findings

Table 1 reports, for comparison, both the OLS linear regressions and the maximum likelihood estimates of the Poisson model for children under age five, for all women age 15 to 49 in Thailand in the 1981 SES.[9] Two empirical specifications are shown, the first allowing for the hypothesized nonconstant returns to scale in family planning, the substitution between public and private program activities, and quadratic effects of women's education and education interacted with age. These added variables are jointly statistically significant. Twice the difference in the log likelihoods for the Poisson model, which is distributed as χ^2, has the value 122, given the assumed density of c_i, suggesting the nonzero effects of these additional variables on fertility. The Poisson parameters can be interpreted as proportional

Table 1
Number of Own Children Less than Age Five
Per Women Age 15 to 49: Thailand 1981 SES a

Explanatory Variables:	Poisson Model (1)	(2)	Ordinary Least Squares (3)	(4)	Sample Statistics b (5)
Intercept	-9.03 (20.4)	-11.0 (27.9)	-2.26 (13.5)	-2.69 (18.3)	.375 (.644)
Age woman (years)	.496 (3106)	.556 (36.4)	.134 (30.4)	.142 (36.9)	29.1 (9.83)
Age^2 (x10-2)	-.891 (3306)	-.939 (35.3)	-.230 (35.9)	-.237 (38.9)	9.42 (.614)
Education woman (years)	-.231 (8.85)	-.0850 (15.9)	-.0357 (4.10)	-.0296 (16.2)	5.26 (3.37)
$Education^2$ ($x10^{-2}$)	-.459 (4.08)	-	-.113 (3.08)	-	.390 (.507)
Age*Education ($x10^{-2}$)	.736 (9.58)	-	.0845 (4.12)	-	1.42 (.932)
Household monthly consumption per adult (log)c	.406 (7.66)	.453 (8.75)	.198 (8.96)	.210 (9.52)	7.00 (.269)
Family Planning Activity: (Baht per woman 1975)					
Public	-.109 (4.67)	-.0230 (4.21)	-.0496 (4.99)	-.00857 (4.21)	9.28 (2.72)
$Public^2$ ($x10^{-1}$)	.289 (2.79)	-	1.38 (3.29)	-	.935 (.582)
Private	-.671 (3.94)	-.142 (4.59)	-.308 (4.39)	-.0576 (4.72)	.619 (.445)
$Private^2$ ($x10^{-1}$)	.841 (.76)	-	.505 (1.12)	-	.0581 (.0482)
Public*Private ($x10^{-1}$)	.515 (4.58)	-	.226 4.99	-	.588 (.464)
Resident in municipal area	-.154 (4.48)	-.167 (4.88)	-.0568 (4.47)	-.0608 (4.84)	.354 (.478)
R^2	-	-	.1502	.1457	
Log likelihood	-8972.1	-9032.9	-	-	
Sample size	12,799				

[a] The absolute value of the asymptotic t ratio for the parameter estimates of the Poisson model are consistent even when there are certain types of specification error due, perhaps, to omitted explanatory variables that are independent of the regressors (Gourieroux et al. 1984), by use of pseudo maximum likelihood methods. Absolute values of t ratios reported in parentheses beneath OLS coefficients.

[b] Means are reported and standard deviations in parentheses beneath the means. The first entry, in the row with intercepts, is the mean and standard deviation of the dependent variable. Divide the OLS coefficient by the mean of the dependent variable to obtain the proportional effect on expected fertility that is estimated in the Poisson specification. For example, the OLS coefficient on public family planning activity (-.00857) in Column (4) divided by .375 equals -.0229 or approximately the Poisson estimate of the proportional effect of a year of education of -.0230.

[c] Endogenous and estimated by instrumental variables where the instruments are the husband's age, age squared, education, household unearned income, amount of unirrigated and irrigated land owned by household.

effects on fertility, i.e., the proportional change in fertility associated with a unit change in the explanatory variable. They can thus be compared to the OLS coefficients at the mean of the sample by dividing the OLS coefficient by the mean level of fertility, i.e., .375. Column 5 reports the means and standard deviations of the dependent and explanatory variables.

Age affects fertility nonlinearly over this range of ages. According to estimates in col. (1), a maximum fertility rate in the preceding five years occurs for women with no education at age 28, and later for more educated women. If the woman's education is one year higher than average, her fertility rate at the sample means is 6.5 percent lower (or 8.5 percent lower in the noninteractive specification (col. (2)). Thus a standard deviation increase in education (3.4 years) is associated with a 22 percent decline in fertility. At higher levels of education, the proportional effect on fertility per year completed becomes larger. A standard deviation increase (27 percent) in predicted household expenditures per adult increases fertility by 11 percent. There was no evidence that the quadratic or cubic specifications of the income effect on fertility fit the data better than the single parameter log income specification reported here. These data, therefore, do not confirm that at higher income levels, proportionate increases in permanent income exert diminishing effects on fertility.

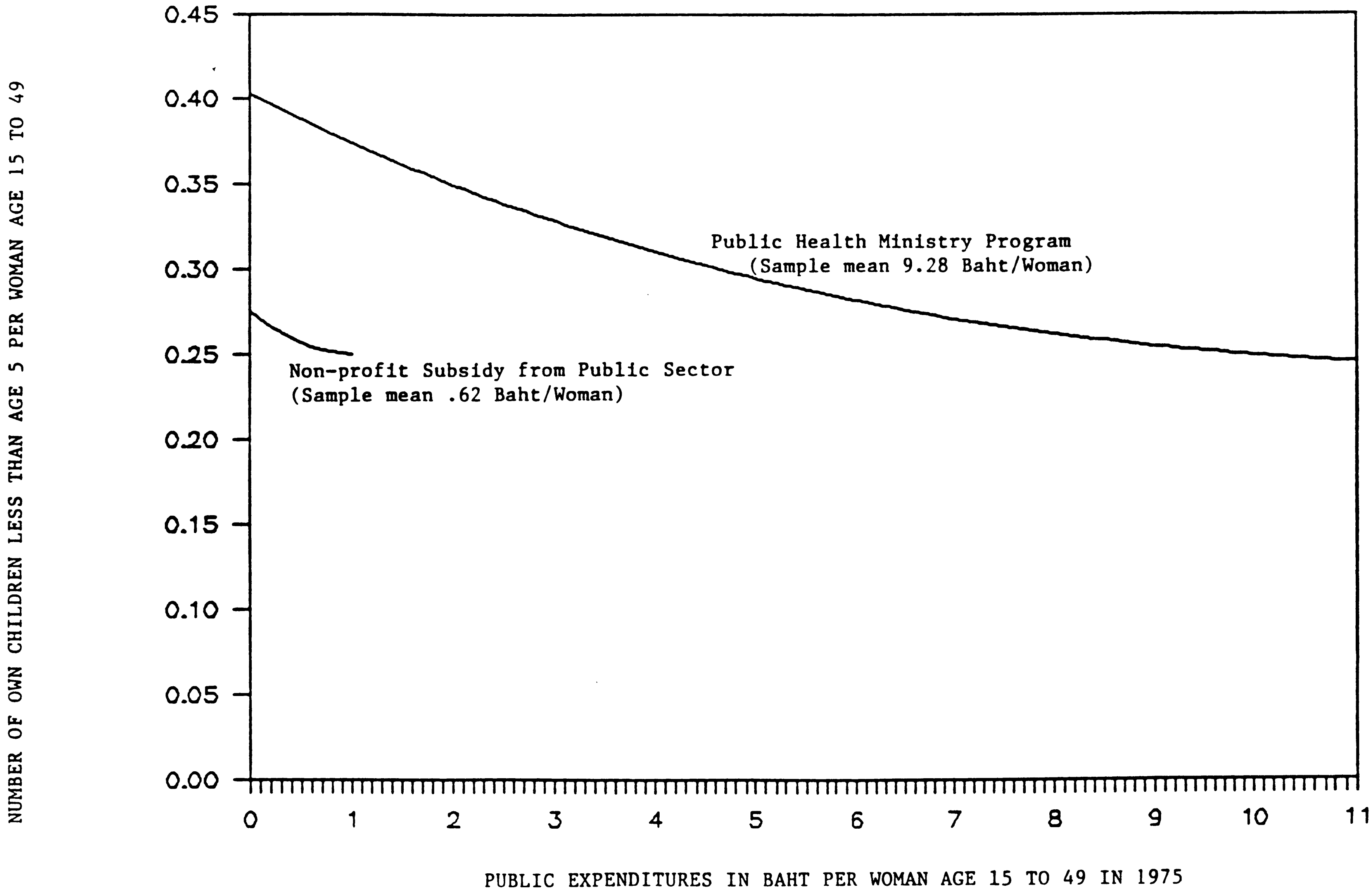
The two family planning coefficients in col. (2) are jointly statistically significant (the $\chi^2 = 24.2$ with two degrees of freedom), and both also are individually significant. A baht (the local currency equivalent to US¢5) spent in 1975 in the clinic-based public family planning program for every childbearing aged woman in a province is associated, on average, with a 2.30 percent reduction in fertility. The same outlay is associated with a 14.2 percent reduction in fertility if it were allocated to the smaller nonprofit family planning program.

Other benefits may, of course, accrue from expenditures on these programs that are not measured here, and these associations undoubtedly include also the effect of program expenditures in neighboring years.

The quadratic approximation specification in col. (1) confirms the further hypotheses that the public and private family planning programs are substitutes for one another, (i.e., positive interaction effects) and that expenditures on the larger public program are subject to diminishing returns to scale. This pattern of diminishing proportionate returns to scale is not statistically significant at the lower levels of expenditure on the nonprofit program in 1975 (.62 versus 9.28 bahts per woman in the private and public programs, respectively). The relationship between family planning expenditures in 1975 and fertility in 1976-81 is simulated in Figure 2, using the preferred estimates of the interactive Poisson model specification col. (1) in Table 1. The slope of these fertility functions, evaluated at sample means, represents the marginal productivity of a baht spent per woman of childbearing age in a province on either the public or private program. The marginal product in the public program (of increasing expenditures from 9 to 10 bahts) is to reduce the fertility level by 2.4 percent compared to that in the private program (of an increase from .6 to 1.6 baht) which would reduce fertility by 8.9 percent. Thus, the payoff to allocating government resources appears to be almost four times as large as the current margin in the private nonprofit versus the public sector. This is because expansion of the public program has already diminished returns, whereas the small subsidy to private nonprofit programs has not yet expanded sufficiently to reduce the returns, though the data suggest this will occur with further program expansion in the late 1970's and 1980's.

FIGURE 2

POISSON MODEL SIMULATION OF ASSOCIATION BETWEEN PROVINCE FAMILY PLANNING EXPENDITURES AND FERTILITY



Source: Table 1, Col 1 and 5

The difference between the predicted level of fertility conditional on zero program expenditures and the observed mean level of fertility is one way to approximate the total effect of the 1975 program on Thai fertility in 1976-81. According to the interactive Poisson (1) specification of the program-production relationship, the public program is responsible for fertility being .25 rather than .40 (intercept in Figure 2); the health ministry's activities thus account for a decline in Thai fertility of 37 percent. The private program accounts for an additional 7 percent decline. According to such a simulation, the absence of both programs would have increased Thai fertility 44 percent above that level actually observed in the 1981 survey.

It should be emphasized, however, that these estimates of the total effect of the program are subject to much greater uncertainty than are the comparisons of the marginal product of the various program expenditures. This is because the slope of the fertility function at the sample means is relatively precisely estimated, and should be a locally unbiased second-order approximation of the marginal effect of program activity (Fuss and McFadden, 1978). The zero program counterfactual case is, however, an extrapolation of the model outside of the range of sample observations. It is consequently sensitive to choice of functional form (i.e., Poisson and linear), and this choice is relatively arbitrary because it is not based on any economic theory or technical knowledge of the operation of family planning programs in general or the particular program in Thailand. Nonetheless, the rough empirical magnitudes suggest that the public family planning program has played a major role in facilitating the 50 percent decline of fertility in Thailand up to 1981.

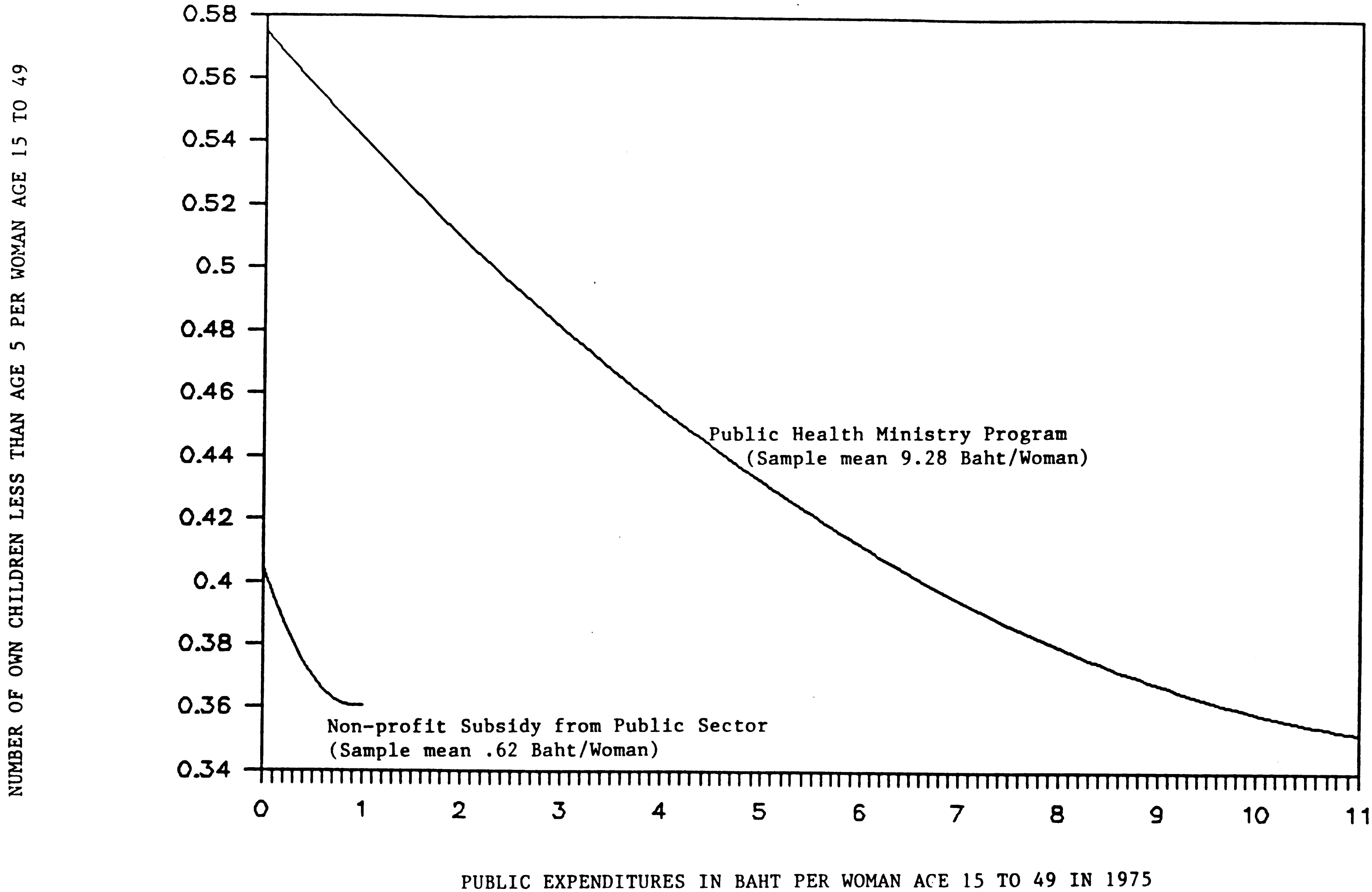
A more conservative approach to estimating the program's total effect on fertility is to attribute the marginal product of the factors (i.e., expenditures on the programs) to inframarginal outlays. This procedure is used where other variable or fixed factors behave as inputs in a production process, and returns to all factors are distributed competitively. In this approach the public family planning program is responsible for a reduction in fertility of 22 percent (i.e., .024* 9.28 Baht/woman), and the private non-profit program subsidy for a 5.5 percent reduction (i.e., .089* .62). The total fertility decline attributable to the two programs, according to this calculation, is still substantial, or more than half of the nationally recorded decline.

Because of the positive effect on fertility of the interaction of private and public program expenditures, inputs to the private program are enhanced by a factor of three if they are concentrated in those areas that receive only half the national average level of public family planning outlays. Conversely, the public family planning program's effectiveness is increased 30 percent when it is concentrated on the areas where the private program is at half its average national strength. The same general patterns and magnitudes can be derived from the linear OLS regressions in column (2) of Table 1, but simulations in Figure 3 illustrate the sensitivity of estimates of total program impact to this change in functional form.

Table 2 repeats the estimations for fertility measured over the longer time horizon, using the number of surviving coresidential own children under ten years of age. Mortality and separation of child from mother is a more serious problem in this case, but one is also closer to a completed span of fertility for many women. Almost twice as much of the variance in this dependent variable is explained by the linear regression model as in that reported in Table 1. The estimated proportionate response of fertility to the covariates in the

FIGURE 3

LINEAR REGRESSION MODEL SIMULATION OF ASSOCIATION BETWEEN PROVINCE FAMILY PLANNING EXPENDITURES AND FERTILITY



Source: Table 1, Col 1 and 5

Table 2
Number of Own Children Less than Age Ten
Per Women Age 15 to 49: Thailand 1981 SES a

Explanatory Variables:	Poisson Model		Ordinary Least Squares		Sample Statistics b
	(1)	(2)	(3)	(4)	(5)
Intercept	-8.86 (27.1)	-10.6 (37.2)	-5.54 (21.34)	-5.84 (25.5)	.806 (1.07)
Age woman (years)	.494 (44.7)	.544 (52.1)	.316 (46.0)	.309 (51.7)	29.1 (9.83)
Age^2 ($x10^{-2}$)	-.783 (46.5)	-.825 (49.0)	-.483 (48.6)	-.477 (50.3)	9.42 (.614)
Education woman (years)	-.199 (9.51)	-.0779 (19.4)	.00992 (.73)	-.0528 (18.6)	5.26 (3.37)
$Education^2$ ($x10^{-2}$)	-.488 (5.90)		-.347 (5.33)	-	.390 (.507)
Age*Education ($x10^{-2}$)	.601 (10.5)		.0386 (1.21)	-	1.42 (.932)
Household monthly consumption per adult (log) c	.377 (9.73)	.402 (10.5)	.395 (11.5)	.394 (11.5)	7.00 (.269)
Family Planning Activity: (Baht per woman 1975)					
Public	-.121 (7.06)	-.0207 (5.05)	-.115 (7.46)	-.0160 (5.06)	9.28 (2.72)
$Public^2$ ($x10^{-1}$)	.0359 (4.77)	-	.0347 (5.33)	-	.935 (.582)
Private	-.790 (6.15)	-.203 (8.86)	-.741 (6.79)	.164 (8.62)	.619 (.445)
$Private^2$ ($x10^{-1}$)	.154 (1.73)	-	.149 (2.12)	-	.0581 (.0482)
Public*Private ($x10^{-1}$)	.500 (6.09)	-	.480 (6.83)	-	.588 (.464)
Resident in municipal area	-.271 (10.7)	-.289 (11.5)	-.226 (11.5)	-.228 (11.7)	.354 (.478)
R^2 -	-	-	.2598	.2541	-
Log likelihood	-13475.5	-13592.7	-	-	-
Sample size	12,799				

a The absolute value of the asymptotic t ratio for the parameter estimates of the Poisson model are consistent even when there are certain types of specification error due, perhaps, to omitted explanatory variables that are independent of the regressors (Gourieroux et al. 1984), by use of pseudo maximum likelihood methods. Absolute values of t ratios reported in parentheses beneath OLS coefficients.

b Means are reported and standard deviations in parentheses beneath the means. The first entry, in the row with intercepts, is the mean and standard deviation of the dependent variable. Divide the OLS coefficient by the mean of the dependent variable to obtain the proportional effect on expected fertility that is estimated in the Poisson specification. For example, the OLS coefficient on education in Column (4) divided by .375 equals -.0816 or approximately the Poisson estimate of the proportional effect of a year of education of -.0831.

c Endogenous and estimated by instrumental variables where the instruments are the husband's age, age squared, education, household unearned income, amount of unirrigated and irrigated land owned by household.

OLS specification are approximately the same as implied by the Poisson model, with the exception of interaction and quadratic variables. One change may be noted with variation in the time horizon; the effect of municipal residence in reducing fertility is more important over the ten-year horizon than over the last five years, confirming that rural-urban fertility differences were closing during the decade before 1981 (Knodel et al., 1987).

Many of the nonlinear terms are less precisely measured in the disaggregated age groups as would be expected, particular those in groups for which the variability is not restricted. Consequently, the simpler specification of the Poisson model without interactions is reported in Table 3. The fertility differentials by education are larger for the 20-29 and 40-49 age cohorts than for the 30-39 cohort, whereas the differences in fertility by income are not statistically significant in the oldest age group. The family planning program appears to have the same pattern of effects on the fertility of women at all ages, but they cease to be statistically significant among women age 40-49.

Interactions between the characteristics of the individual and provincial family planning expenditures in the fertility model are reported in Table 4. The woman's education and local expenditures on private family planning appear to be substitutes in lowering fertility (col. 1). Education of women may reduce the information costs of understanding and adopting new birth control technologies. Public sector programs, however, do not disproportionately benefit less-educated women; they appear to reduce the fertility of better educated women more than they do the fertility of the less educated women. This pattern is unanticipated unless the public sector activities and expenditures are oriented to serve the better-educated, urban middle class with their increasing demand in this period for relatively-

Table 3
Number of Own Children Less than Age Five
Per Woman, by Age of Woman: Thailand 1981 SES [a]
(Poisson Model)

Explanatory Variables:	Woman's Age 20-29 (1)	30-39 (2)	40-49 (3)
Intercept	-16.9	-.0463	-26.7
	(10.2)	(.01)	(1.78)
Age Woman (in years)	.959	-.0967	1.41
	(7.30)	(0.42)	(2.10)
Age Squared ($x\ 10^{-2}$)	-.185	.0277	-1.78
	(6.96)	(.06)	(2.33)
Education woman (in years)	-.108	-.0233	-.0526
	(16.9)	(4.76)	(2.26)
Household Monthly Consumption per adult (log)[b]	.726	.379	.302
	(11.2)	(4.77)	(1.42)
Family Planning Activity (Bahts per woman in 1975)			
Public	-.0228	-.0154	-.0235
	(3.26)	(1.64)	(1.13)
Private	-.152	-.120	-.0574
	(3.92)	(2.16)	(.49)
Resident in Municipal Area	-.219	-.0873	-.497
	(4.89)	(1.50)	(3.47)
Log likelihood	-4281.6	-2804.9	-1120.6
Sample Size	4490	3159	2485
Dependent Variable	.604	.482	.163
(standard deviation)	(.757)	(.682)	(.423)

[a] The absolute value of the asymptotic t ratio for the parameter estimates of the Poisson model are consistent even when there are certain types of specification error due, perhaps, to omitted explanatory variables that are independent of the regressors (Gourieroux et al. 1984), by use of pseudo maximum likelihood methods.

[b] Endogenous and estimated by instrumental variables where the instruments are the husband's age, age squared, education, household unearned income, amount of unirrigated and irrigated land owned by household.

Table 4

Number of Own Children Less than Age Five Per Woman Age 15 to 49: Family Planning Program Interactions [a]

Explanatory Variables:	Education of Women (1)	Household Consumption (2)	Education and Consumption (3)	Municipal Residence (4)	North-East Region (5)
Program Interactions:					
Public*education ($x10^{-2}$)	-.403 (1.96		-.466 (2.28)		
Private*education ($x10^{-2}$)	1.18 (.92)		.0871 (.07)		
Public*household consumption [b]		.00476 (.26)	0166 (.89)		
Private*household consumption b		.352 (3.03)	.349 (3.07)		
Public*municipal				.00348 (.28)	
Private*municipal				.198 (2.64)	
Public*Northeast					-.0280 (2.14)
Private*Northeast					.118 (1.43)
Northeast region					.395 (3.46)
Other variables reported in Table 1					
Log Likelihood	-8969.2	-8966.6	-8964.4	-8968.3	-8956.1
χ^2 Likelihood ratio statistical	5.8	11.0	15.4	7.6	32.0
significance level over Col. (1) estimates Table 1	.10	.005	.005	.05	.005

[a] Poisson model where all the other explanatory variables included in the interactive specification in column 1 in Table 1 are also included but not reported here for brevity.

[b] The logarithm of household consumption per adult is treated as endogenous and thus this interaction variable is endogenous and estimated by instrumental variables where the instruments are the husband's age, age squared, education, household unearned income, amount of unirrigated and irrigated land owned by household.

costly surgical sterilization procedures. The permanent income (i.e., household consumption) interactions are jointly statistically significant at a $p < .005$ level in cols. 2 and 3. In this case the lower-income household gains more than the richer in terms of the number of births averted by a given program expenditure. The distribution of the fertility-averting benefits of the family planning program are more clearly concentrated among the poor in the case of the private nonprofit program than in the government-run program.

Statistically the private program has a significantly larger negative effect on fertility in rural (nonmunicipal) areas than in the urban (municipal) areas, as seen from col. 4 of Table 4. The Northeast region of Thailand is the poorest; it had experienced the smallest relative decline in fertility by 1981. Among women living in the Northeast, the public sector appears to have exerted a particularly strong effect in lowering fertility (col. 5) or an extra 3 percent decline in fertility per baht per woman. The impact of private sector activity in this region is less notable than elsewhere in Thailand.

Both private and public family planning programs are thus associated with lower fertility, and these behavioral changes are generally concentrated among the more disadvantaged classes of Thailand, measured by household income, rural residence, or residence in the impoverished Northeast region of Thailand. Though statistically weak, these program interactions suggest a given budget for private non-profit family planning expenditures would have been more effective if it had been more concentrated in the provinces that have heretofore received the lowest levels of expenditure per woman of childbearing age. Moreover, somewhat more public resources at the margin could be productively used after 1975 in the private nonprofit organizations, rather than in the expansion of the programs of the Public Health Ministry. In practice, such a reallocation might be accomplished if fees are used to finance a share of the costs of the Health Ministry's urban program in the richer regions of Thailand. Because the program related fertility reduction of women is greater among those living in low-income (or consumption) households, price-discrimination in favor of these poorer households could increase the fertility reducing effect of a given level of government program expenditures.[10] To avoid the administrative cost of collecting user-fees on the basis of difficult to monitor household consumption levels, regional prices might be varied.

To explore this possibility further, Table 5 reports the children 0-4 linear regressions separately for the urban, sanitary district (suburban), and rural populations. Comparisons of effects confirm that the impact of women's education, permanent income, or family planning is not limited in one strata of the society. Indeed, the empirical literature on fertility would have led to the expectation that the positive permanent income effect on fertility would have diminished among higher-income urban populations, but the estimated income elasticity is twice as large in municipal as in rural or suburban Thai populations. But private nonprofit family planning outlays are particularly effective outside of the municipal areas where the commercial market for contraceptives may be less well established.

Community's education and health programs and unobserved community constraints relevant to fertility may be correlated with the observed individual couple's characteristics, such as household consumption, mother's education, and age. Consequently, the effects of these individual structural variables are reestimated in col. 4 of Table 5 within communities defined as the 279 sampling clusters in the survey. These estimates, as with the urban,

Table 5
Number of Own Children Less than Age Five
Per Woman Age 15 to 49: Stratified by Region [a]

Explanatory Variables:	Municipal Urban (1)	Sanitary Districts (2)	Rural Villages (3)	Within Sample Clusters [b] (4)
Intercept	-5.71 (16.6)	-2.97 (8.11)	-3.03 (9.98)	.0042 (.81)
Age Woman	.105 (14.1)	.118 (12.2)	.153 (26.2)	.134 (36.5)
Age Squared ($x\ 10^{-2}$)	-.182 (15.5)	-.201 (13.2)	-.254 (27.5)	-.222 (38.5)
Education Woman	-.0299 (11.7)	-.0248 (5.22)	-.0305 (8.21)	-.0284 (15.3)
Household Monthly Consumption per Adult(log) [c]	.672 (13.4)	.290 (5.15)	.247 (5.11)	.250 (11.1)
Family Planning Activity per Woman in 1975				
Public	-.00787 (1.99)	-.00742 (1.41)	-.00893 (3.11)	--
Private	.00555 (.81)	-.0872 (2.75)	-.0914 (5.12)	--
R^2	.1370	.1326	.1485	.1400
Sample Size	4491	1954	6354	12799

a Ordinary least squares with absolute values of t ratios reported in parentheses beneath coefficients

b All variables expressed as deviations from means within each of the 279 sample clusters. Estimates of household variable effects on fertility estimated within clusters are free of bias arising because locality unobserved determinants of fertility that are correlated with household variables analyzed.

c Endogenous and estimated by instrumental variables where the instruments are the husband's age, age squared, education, household unearned income, amount of unirrigated and irrigated land owned by household.

suburban, and rural strata, are similar to the overall estimates based also on the between cluster variation in fertility (Table 1, col. 4). The F test of the statistical improvement gained by including these 279 regional fixed-effects is, however, not statistically significant, nor are the community effects correlated with the individual covariates in the fertility model. Of course, it is not possible to assess whether the family planning program effects are biased by such omitted community effects, because family planning expenditures do not vary within the community clusters.

Prior studies of Thai fertility help assess these results. Kiranandana et al. (1984) estimated the determinants of total fertility rates in 1980 across provinces, including public private family planning expenditures, and GNP per capita in 1975. Although they found that private nonprofit family planning expenditures from the public sector were more strongly related to fertility than public outlays channeled through the larger public health ministry's budget. However, they constructed a single measure of "economic development" in the region that was not related to fertility. They concluded that economic changes were not responsible for much of Thailand's fertility decline. Many fertility studies also find weak relationships between fertility and <u>total</u> family income or a development index. But women's labor supply is determined simultaneously with fertility and is included in total family income and measured GNP per capita. Different sources of family income—those derived from the value of men's time and the value of women's time and non-human capital wealth and natural resources—embody different price and income effects on fertility. These different sources of income exert different effects on fertility in many empirical studies (Schultz, 1997). Treating family income as endogenous, and instrumenting it with exogenous characteristics of male earnings and non earned income helped in this study to estimate the underlying, more stable parameters associated with a woman's value of time proxied by her education, and with male productivity and family nonearned income, including land ownership as proxied by family consumption per adult.

Another study of the allocation of family planning resources in Thailand found an effect of program staff assignments on contraceptive use within the public program (ESCAP, 1986). However, contraceptive usage in a family planning program is not a satisfactory indicator of program impact, because substitution of contraceptive demand may occur from private market suppliers to the subsidized public sector, overstating the program's effect on new users and hence its net effect on fertility (Schultz, 1988). Nor is it reasonable to specify "contraceptive" use as an exogenous determinant of fertility in a block recursive multiple regression system, as proposed in the ESCAP study. To evaluate properly how contraceptive use affects fertility, the practice of contraception must itself be explained or endogenized. Moreover, multicollinearity among aggregate explanatory variables at the provincial level, as analyzed in both of these studies, may make it difficult to estimate suitable models of fertility from only provincial aggregates. If the micro relationships are nonlinear and interactive, as suggested here, the provincial aggregate observations on explanatory variables are not suitable population "averages" for the province, but instead more complex weighted functions of other moments are needed. The individual level survey data is expected to yield more flexible nonlinear relationships and test more detailed hypotheses regarding who benefits from public sector expenditures on family planning. The personal distribution of public program benefits is almost as important as the total magnitude of program effects on fertility in justifying public sector involvement in population programs.

VI. Conclusions

Survey information on the income, education, and fertility of couples, in conjunction with data on local expenditures on two forms of family planning activities were examined for Thailand, a country in which total fertility rates declined by nearly one-half in the two decades before the 1981 survey. For a country in which only 15 percent of the population lived in municipal areas and per capita income was US $380 in 1976, this demographic transition is notable.[11] The impact of two family planning program strategies was assessed to determine how this type of public service affected the fertility of different segments in Thai society.

A Poisson model of the number of surviving coresidential own-children per woman was fit to these data by robust maximum likelihood methods. These estimates do not qualitatively differ from linear ordinary least squares estimates, but they offer a somewhat more plausible and precise representation of the discrete nature of recent fertility rates for individual women. For evaluating family program effects at the individual level, the own-children measure of women's recent fertility is an empirical compromise imposed by the limitations of data, for it represents a hybrid measure of "surviving" fertility that has no exact counterpart in lifetime behavioral or biological models. But for policy evaluation to proceed in a timely fashion, it is generally not practical to wait until lifetime cumulative fertility of cohorts can be linked to past program expenditures.

The estimates of the Poisson model for surviving fertility rates of women in Thailand describe household economic and locational variables that are significantly associated with fertility. Education differentials are still large, though, as noted in a later survey, closing

(Chamratrithirong et al., 1987). Household income appears to relax the economic constraint on having more children.[12] Although incomes are rising rapidly in Thailand, other factors making children more costly or less valued have offset these gains in income. The increasing education of women is one development that has strongly countered the effect of increasing male and nonearned incomes on fertility.

A major factor in facilitating fertility decline has been the Thai government's family planning programs. One program is under the Public Health Ministry, employing physicians and nurses in public health clinics and hospitals. It is a costly program, but has had a major impact on fertility, particularly for those living in municipal areas where most of these facilities are located, and particularly through the provision of sterilization which had become, by the 1980's, the most frequently used contraceptive method.

The private nonprofit family planning associations are subsidized by the government, but these subsidies in 1975 were only a small fraction of the public resources that were allocated to family planning programs operated directly by the Health Ministry. Nonetheless, governmental subsidies to the private nonprofit program have had a particularly large relative effect on fertility rates in those provinces where the public health program is small. These patterns of program effectiveness and substitution are interpreted with the aid of a simple economic model. They have interesting implications for the design and management of family planning programs in low-income countries such as Thailand, where private market distribution channels for some major forms of contraception, such as the pill, condom, and injectable, have developed rapidly in the wake of the demographic transition, and private nonprofit family planning associations have provided innovative leadership in serving specific segments of the Thai population.

Table A-1

Number of Men and Women in Working Sample From 1981 SES

	Persons Over Age 14	Reporting Wage
Women total	13871	4145
Women without matched spouse	7109	
Men total	12,242	6041
Men without matched spouse	5480	
Women and men (matched couples)	6762	
Total working sample	19,351	10,186
Unattributable children or spouses in household preclude confident match	325 (1.65% of total)	
Total file	19,676	

Table A-2

Estimates of Private Rates of Return to Schooling by Sex, Using Alternative Specifications and Estimation Methods: 1981 SES

Specification Sex Estimation Method	Private Rates of Return by School Level:[a] Primary	Secondary	Higher	Average
1. Wage Function Excludes Regional Variables				
Females (4145/13871)				
OLS	.13	.37	.11	.27
	(7.58)	(30.8)	(3.67)	(43.3)
Heckman Two Step	.17	.30	-.11	.22
	(10.3)	(.24)	(3.16)	(32.5)
Maximum Likelihood	.15	.40	-.03	.20
	(8.06)	(20.9)	(.45)	(19.3)
Males (6041/12242)				
OLS	.25	.33	.12	.25
	(15.6)	(30.4)	(4.19)	(46.7)
Heckman Two Step	.22	.30	-.021	.25
	(14.2)	(29.1)	(.72)	(43.1)
Maximum Likelihood	.19	.30	.020	.23
	(9.47)	(18.6)	(.40)	(27.1)
2. Wage Function Includes Regional Variables				
Females				
OLS	.066	.29	.069	.19
	(4.40)	(27.0)	(2.78)	(34.0)
Heckman Two Step	.097	.23	-.10	.16
	(6.53)	(20.0)	(3.26)	(24.1)
Maximum likelihood	.11	.22	-.13	.14
	(5.56)	(13.5)	(2.41)	(15.3)
Male				
OLS	.13	.24	.078	.19
	(9.46)	(25.7)	(3.38)	(37.1)
Heckman Two Step	.12	.23	.006	.17
	(8.89)	(24.4)	(.24)	(29.6)
Maximum Likelihood	.12	.23	.058	.17
	(7.54)	(17.7)	(.17)	(24.0)

Notes:
[a]The private rate of return is approximated by the coefficient estimated on the years of schooling variable in a logarithmic wage rate regression that also includes as regressors the individual's years of postschooling experience and experience squared. The sample includes all persons 15 to 49 in the 1981 SES.

Table A-3

Sample Composition	Sample Size	Mean	Variance	Mean Square Error of Regression Model[a]
All Women				
Age 15 to 49	12620	.376	.416	.354
Age 15 to 19	2632	.062	.072	.063
Age 20 to 29	4422	.604	.574	.507
Age 30 to 39	3120	.485	.467	.449
Age 40 to 49	2446	.165	.181	.175

[a]Regression model is specified as eq. (3) in Table 1.

Statistical Appendix

The 1981 SocioEconomic Survey sampled 30 amphoes (districts) in each of the four regions of Thailand outside of Bangkok, and 30 districts in Bangkok. Four strata were distinguished for district sampling rates: the central special region (1/300), the municipal areas (1/250), the sanitary districts (here called suburban) (1/500), and the rural remainder (1/100). Each household was interviewed and reinterviewed when substantial discrepancies between recorded income and expenditures were uncovered. The National Statistical Office (NSO) planned and implemented the survey, processed and coded the data for the publicly released file.

A distinct feature of the 1981 SES is that it treated the sub-household with its own income as separate insofar as enumeration, whereas in the past, (i.e., 1975 SES and 1970 SES) multigenerational households in which each generation has some independent income were combined. More small households are therefore evident in the 1981 SES than in the 1980 Census. Unless the broader household unit is pooling resources, the nuclear family focus should be helpful in isolating the economic resources and expenditures that are most immediately relevant to the couples' reproductive decisions. Given the limitations of survey codes for the relationship of each individual to the head of household, the redefinition of the household to be the nuclear (not intergenerational) family is helpful in matching up spouses and allocating children to their mothers.

Out of 19,676 men and women in the sample over age 15 reporting the core demographic questions, 325 individuals could not be confidently matched either to children who might possibly have been their own children or to a spouse in the household. Six particular linkage problems are distinguishable. Ninety-two men were married but linked to the head as only "other relative" and thus not a spouse, child or child-in-law. One hundred sixty-five married women fell into this category as well. Ten men were heads of their households but had more than one spouse in the household, making it impossible to allocate children of the head to their biological mothers. These apparently polygamous unions are not uncommon in the southern Muslim region of Thailand, but they appear to normally reside in separate households, and presumably present fewer matching problems for my analysis. Twenty-two men were children of heads, but there were more than one female child-in-law of the head in the household to whom the men might have been married. Conversely, 36 female children of the head have multiple possible spouses related to the head as "children-in-law." The small attrition rate from the sample of 1.65 percent is probably due to the emphasis given in the 1981 SES to define the interview unit as the nuclear, self-supporting household and not the traditional, extended, coresidential household (Chiswick, 1987). Table A-1 provides a breakdown of the individuals from the working sample of the 1981 SES by sex and whether their spouse was enumerated in the household. Table A-1 also shows that three-tenths of the women reported a wage or salary and hours worked, whereas almost one-half of the men reported these data. These data were required to compute an hourly wage rate, and infer how schooling is related to hourly wages, after correcting for sample selection bias (Schultz, 1993).

Notes

1. The complexity of dynamics and life cycle context in which fertility occurs suggests the desirability of measuring how public family planning helps couples have the children they want, when they want them, over their lifetime. In this paper, the more limited question of how period fertility rates are affected by the program is examined (Schultz, 1988). There is also a large literature that discusses the problems of evaluating program effects on human behavior and achievements, such as the literature on job training programs in the U.S. These statistical and conceptual problems are presented by Heckman and Hotz (1987).

2. These simple wage comparisons suggest an average private rate of return to education in Thailand for both men and women of between 25 and 27 percent using a Mincerian (1974) wage function specification. These calculations, however, do not take account of the possibly unrepresentative character of wage earners. The estimates are then corrected for this potential sample selection bias, identifying the probit selection rule by the inclusion of irrigated and unirrigated land owned by the household and nonearned income of the household and four regional dummies: Bangkok, municipality, suburban (sanitary district), and northeast region. Other identifying restrictions were used including, in the wage selection equation, the self-employment status of the spouse and whether a spouse is present in the household, with little change in the estimates. The four regional dummy variables are also included in the wage function itself, thus eliminating the returns to education that are associated with migration from rural to urban higher wage areas (Schultz, 1993).

3. See later discussion of the 1981 SES new concept of the nuclear family. See Chiswick, 1987.

4. The expenditure data include imputations for the value of owner occupied housing, payment of wages in kind, and home produced and consumed food and produce. See C. Paxson (1988) for a study of savings rates derived from this survey.

5. The zero order correlation of the count of own coresidential children age 0 to 4 and the count of such children 10 to 14 is -.03 in the full sample examined here of women aged 15 to 49.

6. See Statistical Appendix text and, specifically, Table A-1.

7. Others have developed regional estimate of fertility by analogous methods in Korea, China, and Thailand. See, for example, the work of Arnold et at., 1985, based on the 1980 Census of Thailand.

8. A fixed-effect Poisson model is a method used to eliminate the possible correlation between omitted, but time invariant, variables that will tend to be correlated with the observed X_{it} . In this case, it is essential that one observe the c_i and x_i over t time periods when the ϵ_{it} will have a region-specific fixed effect representing the omitted variables. In the case at hand, child health programs may vary interregionally along with family planning, but be fixed

within a region over time (see, for example, Hausman et al., 1984 and Portney and Mullahy, 1986).

9. Serial correlation over time in the unexplained error or propensity of some couples to have birth might occur, particularly in a high fertility population, where individual variation in fecundity would exhibit itself in the unregulated pace of conceptions across couples. This could possibly be dealt with by using a fixed-effect model (Hausman et al., 1982; Hausman et al., 1984; Portney and Mullahy, 1986), and breaking down the count of children into a series of age-of-children intervals for each woman, i.e., age 1-3, 4-6, 7-9, 10-12, etc. and thereby identifying the common component across intervals for a couple as their fixed effect. Of course, the actual frequency of births will reflect the combined influence of both the serially correlated biological supply capacity (biological heterogeneity or fecundity) and the behavioral demand regulated response (Rosenzweig and Schultz, 1985). Separating these two sources of reproductive outcomes would require information on the history of contraceptive behavior for each couple, a full fertility history, and a time series on local family planning activity, none of which is available to the author for Thailand.

10. Taiwan illustrates a country with an early active family planning program starting in the 1960s which has recently retrenched its subsidies (Schultz, 1988). For example, starting in 1990 sterilization operations were subsidized only for low-income families and those with special health problems. A new family planning program directive in 1995 restricted the subsidy for the insertion of IUDs, the most popular method after sterilization, to one among seven varieties and for only indigent women and special health groups. As a consequence, public program IUD insertions decreased 81 percent from 103,367 in FY 1994 to 19,626 in FY 1995. The crude birth rate incidentally increased 1.2 percent in calendar 1995 over 1994 from 15.3 to 15.5 per thousand persons (Taiwan Provincial Institute of Family Planning, 1996; Taiwan Ministry of Interior, 1996).

11. The likelihood function is

$$L = \sum_i \sum_t \left(c_{it} x_{it} \beta - e^{x_{it}\beta} - \ell n \, (c_{it}!) \right)$$

and the derivatives with respect to β are

$$\frac{\partial L}{\partial \beta} = \sum_i \sum_t \left[x_{it}' (c_{it} - e^{x_{it}\beta}) \right] .$$

12. To my knowledge, only China in the 1970s reduced its birth rate more rapidly at this low an income level. China, however, accomplished its rapid demographic transition by using social pressures that may be viewed as coercive, whereas Thailand has preserved individual reproductive choice and subsidized the provision of voluntary services.

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