

ECONOMIC GROWTH CENTER

YALE UNIVERSITY

Box 1987, Yale Station
27 Hillhouse Avenue
New Haven, Connecticut 06520

CENTER DISCUSSION PAPER NO. 662

THE RELATIONSHIP BETWEEN LOCAL
FAMILY PLANNING EXPENDITURES
AND FERTILITY IN THAILAND,
1976-1981

T. Paul Schultz

Yale University

April 1992

Note: Center Discussion Papers are preliminary materials circulated to stimulate discussions and critical comments. This paper was originally presented at the Annual Meetings of the Population Association of America, March 30, 1989 in Baltimore.

Research for this paper was partially funded by the Women in Development Division, Population and Human Resources Department of the World Bank.

I have benefitted from the discussions regarding this topic with Chalongphob Sussangkarn, Ashakul Teera, Charles Myers, Martha Ainsworth, Angus Deaton, Peter Donaldson, Suganya Hutaserani, John Mullahy and James Walker. Mr. Chalongphob at the Thailand Development Research Institute was of particular help in analyzing the 1981 SES. Paul McGuire provided able programming assistance at Yale.

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 cost effectiveness 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 critically important in accounting for the rapid decline in that country's fertility.

The Relationship between Local Family Planning Expenditures and Fertility in Thailand, 1976-1981

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 on the effects of family planning program expenditures on the decline in fertility (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 does 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 why the marginal effect of public expenditures may change with the scale of programs. Without a consensus on the appropriate methods to use to estimate the cost effectiveness of family planning, this round of reductions in public sector programs is likely to reduce priorities for family planning at a time when population growth continues to rise in Africa toward 3 percent per year and still mounts at 2 percent per year in 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 expenditures on family planning. Recent fertility levels 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 Socioeconomic Survey is designed to obtain accurate data on consumer market expenditures, household consumption of own production, and income in kind. Because consumption proxies permanent expected lifetime income better 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 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 family 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. First, the nature of the policy intervention is described and the modeling issues for empirical study 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 proxies 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 environmental 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 simple birth control supplies and information provided by paramedics such as midwives, who canvas local communities to encourage adoption and use (i.e. outreach), (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 to an informed and receptive public a routine service and source of subsidized contraceptives. 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. Evaluation of the separate segments of the program is consequently often difficult, if not impossible.

Such public sector family planning programs might exhibit several properties: (1) they are eventually subject to diminishing returns to scale as they expand, where scale is defined as the level of local program expenditure per potential beneficiary, or in this case public outlays per woman of childbearing age (Schultz, 1988), (2) 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 is 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, while the rural population is served by the paramedic outreach activities, in which case clinics and outreach may not erode the marginal product to the other program segment; (3) information may either complement or substitute for the effectiveness of the other two distribution networks (Schultz, 1988); (4) 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); and (5) the greater the household's permanent income per adult, the less influential are 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).

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 diminish at higher levels of schooling if education of women affected fertility only to the extent that it increased market wage opportunities and the value of time of women.

The household demand framework 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). Nonearned income should increase the desired number of children, for increases in physical wealth or transfers that add to nonearned income relax the household's budget constraint without increasing the price of children by raising the opportunity cost of the time of either parent in child care (see Supriya, 1982).

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 who constitute the largest group of families in Thailand. Thus, measured household income distinguishes the earnings of all adults in the nuclear family^{6/} and nonearned income from land and other assets. From other studies of fertility determinants in primarily agricultural populations at Thailand's income level, family income from land, other assets and husband's earning are generally associated with higher fertility (Schultz, 1973; Mueller and Short, 1983).

Farmer 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, and this may also be true of other types of self-employed activity in Thailand. Our approach to fertility as a household commodity of demand postulates a responsiveness of fertility to variation in household permanent or lifetime expected income and wages. The household's level of consumption rather than reported income is used as a proxy for permanent lifetime income of the couple. Using the 1981 SES, Paxson (1988) has shown that annual deviations in weather from historic averages explain savings behavior among Thai rice farmers. Because of the relative importance of transitory income among poor self-employed, the household's expenditures per adult per month is specified as the proxy for household permanent income. This variable is treated as endogenous because it includes women's earnings which depend on their labor supply decisions. Measurement error may also favor the analysis of consumption totals rather than income totals of Thai households surveyed in the 1981 SES.

Many questions in this survey were directed to quantifying current, periodic, and durable expenditures, the imputed value of household produced consumption, income in kind, and the rental value of owner occupied housing. This exhaustive enumeration of expenditures and consumption is likely to be

more complete in measuring consumption than the handful of questions devoted to reporting wage and nonwage income of family members. Consequently, the preferred measure of permanent income adopted here is total consumption divided by the number of working age adults (15 to 65) in the household.^{7/} Although this measure of permanent income is a combination of the labor productivity of men and women in the household as well as asset income, if the mother's education is also controlled in the relationship, this household income variable should capture the effect of prenatal sources of family income, while the mother's education proxies for the woman's wage opportunities.

However, this consumption based measure of permanent household income may not be 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 to the estimate of the income effect based on household consumption. Finally, children before they reach age 15 may contribute to family income, 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.^{8/} To avoid these several sources of simultaneous equation bias and bias due to 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 increases in women's wages may be responsible for 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), separately 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 was 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 fertility from 1976 to 1981 more strongly than 1980 family planning expenditures. Empirically, the inclusion of the 1980 family planning program expenditure values, in addition to those reported for 1975, did not improve statistically at the 10 percent level of significance 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 proxy the average of the past several years of family planning expenditures in these programs. Family planning inputs for all years from 1970 to 1980 should be analyzed in future work, employing a smoothed distributed lag specification (Schultz, 1988), but all of these annual series are not 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 analysis of this important issue 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 a third of the SES sample of 12,620 women age 15 to 49 live in urban or municipal areas, whereas nearly half live in rural villages, with the remaining sixth in "suburban" areas called sanitary districts, which are often very similar to the urban municipal areas, but have not been reclassified for political reasons (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 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 modern birth control information and services.

In sum, the recent period rate of fertility of a woman is assumed to be a reduced-form function of local family planning service subsidies, and 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, X_{1p} and X_{2p} represent public expenditure in 1975 on family planning to The Health Ministry and to private nonprofit organizations, respectively, at the local province level, A_w is the woman's age, E_w her completed schooling, I_w her household's consumption per adult (proxy for permanent income), and M_w 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:

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

$$\text{and } d^2F/dX_1dX_2 > 0$$

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

$$(3) \quad dF/dE < 0$$

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

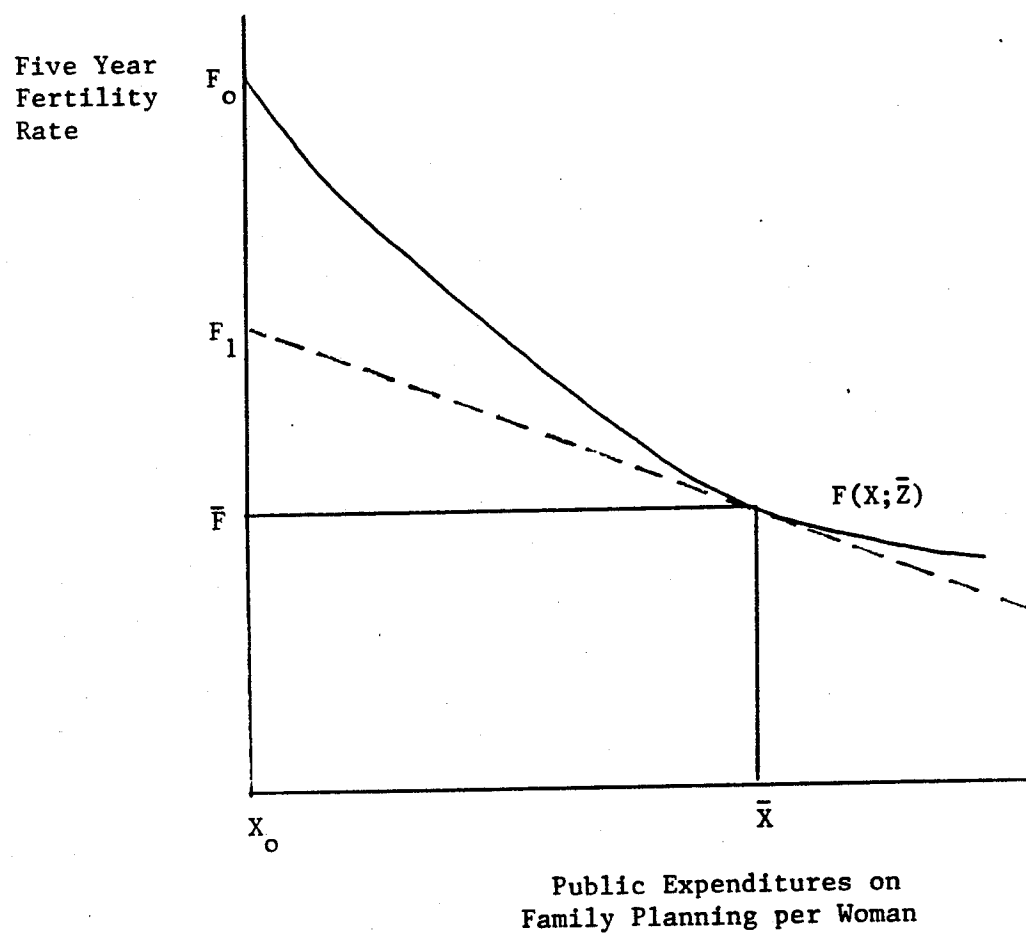
$$(5) \quad dF/dI > 0$$

$$(6) \quad dF/dM < 0$$

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 diminishing returns to scale as hypothesized, the marginal effect of public expenditure at the sample mean will be less than the average

FIGURE 1

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



effect, i.e. $dF/dX = (F_1 - \bar{F})/\bar{X} < (F_0 - \bar{F})/\bar{X}$, 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 in many countries comprehensive reproductive histories, but did not collect even rudimentary household income, wages or expenditure information or matched 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, I have assigned priority to obtaining good household expenditure and income data, but have not been able to disaggregate family planning expenditures below the provincial level, and rely on an indirect measure of recent fertility, derived from the registry of household members identified by their relationship to the head of household.

When vital registration systems are incomplete, as they often are in low income countries, age specific birth rates are indirectly estimated by calculating the frequency of surviving own-children per woman of childbearing age in a census or representative household survey (Lee-Jay 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. In this paper I examine at the level of the individual woman the number of her own

surviving 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 level is infrequently analyzed for several reasons. Two to three percent of the children under 5 years of age cannot be attributed to a mother in the 1981 SES, because household structure information is inadequate for matching or because the mother does not live with her child (see Appendix). A few cases also occur of polygamous unions in the same "nuclear family" that prevent matching a child to her mother. 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. 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.^{9/}

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.^{10/} 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. 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) is thus estimated of 3.03. National estimates of

TFR from other sources are 3.67 in 1975, 3.01 in 1980 and 2.11 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 STFR 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:

$$\text{STFR}_{p76-81} = \frac{.904}{(25.8)} \text{TFR}_{p80} \quad R^2 = .92$$

where $p = 1, \dots, 62$ 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 STFR as an approximation of the recent rate of childbearing in the Thai population.

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 must be specified linking this count of own children and 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:

$$f(c_{it}) = (e^{-\lambda_{it}} \cdot \lambda_{it}^{c_{it}}) / c_{it}! \quad (1)$$

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:

$$\lambda_{it} = e^{x_{it}\beta}, \quad (2)$$

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

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

The maximum likelihood method is used to obtain estimates of β .^{11/} However, a multiplicative error, e_{it} , might arise due to omitted variables of the following form:

$$\lambda_{it} = e^{x_{it}\beta} e^{\epsilon_{it}}. \quad (4)$$

Even when ϵ_{it} is uncorrelated with x_{it} , the maximum likelihood estimates of standard errors may not be consistent.^{12/} These errors are also a potential explanation of the over dispersion problem. A pseudo maximum likelihood procedure proposed by Gourieroux et al. (1984) provides asymptotically unbiased standard errors in this case. These robust standard errors are reported in some economic applications to be substantially larger than those obtained from conventional maximum likelihood methods in the presence of over dispersion (Mullahy, 1986). The corrected robust standard errors are therefore reported below.

Three limitations are apparent in the Poisson analysis of fertility; the first two of which apply 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 does implicitly linear (OLS) regression or some other dynamic representations (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 decade, 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 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, 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-3), whereas the residual 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 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.^{13/} Two empirical specifications shown, the first allowing for the hypothesized nonconstant returns to scale in family planning, the substitution between public and

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

Explanatory Variables:	Poisson Model		Ordinary Least Squares		Sample Statistics b
	(1)	(2)	(3)	(4)	
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.

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 effects on fertility, i.e. the proportional change in fertility associated with a unit change in the explanatory variable. They can be 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.

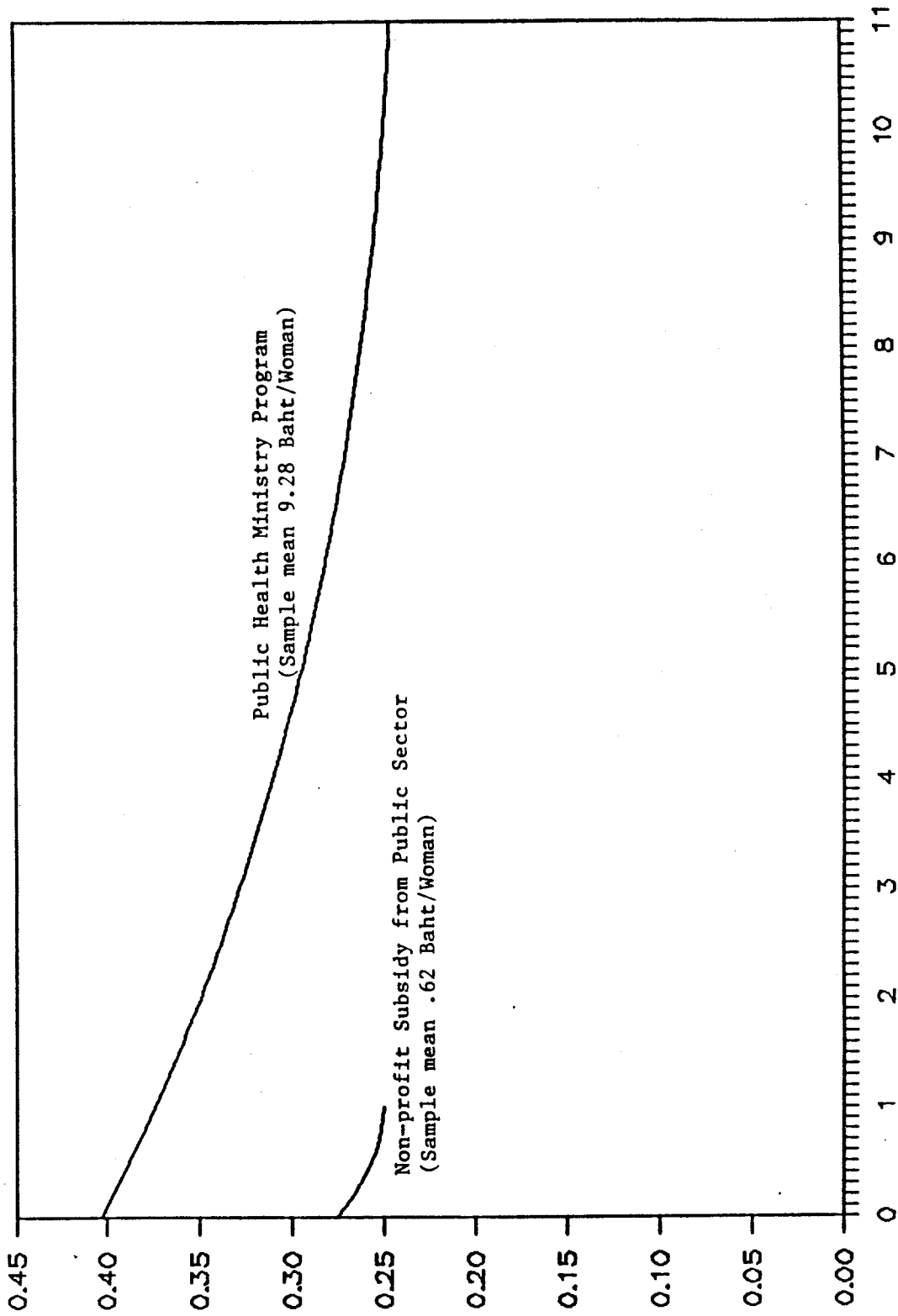
Age affects fertility nonlinearly over this range of ages. According to estimates in col. (1), a maximum fertility rate in the preceding five years occurs for women with no education at age 28, and later for more educated women. If the woman's education is one year higher than average, her fertility 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 expenditures per adult increases fertility by eleven 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 fertility declines with additional permanent income.

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 is allocated to the smaller private 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 private 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 public resources appears to be almost four times as large at the current margin in the private versus the public sector. This is because expansion of the public program has already diminished returns,

Poisson 0-4



PUBLIC EXPENDITURES IN BAHT PER WOMAN AGE 15 TO 49 IN 1975

Source: Table 1, Col 1 and 5

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 program expansion in the late 1970's and 1980's.

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 the two programs would have increased Thai fertility 44 percent above that 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 program expenditure. 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, and it is consequently sensitive to choice of functional form (i.e. Poisson or linear) and this choice is relatively arbitrary because it is not based on any economic theory or knowledge of the operation of family planning programs in general or specifically the 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 factors (i.e. expenditures on the programs) to inframarginal outlays. This procedure is used in economics 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 5.5 percent (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.

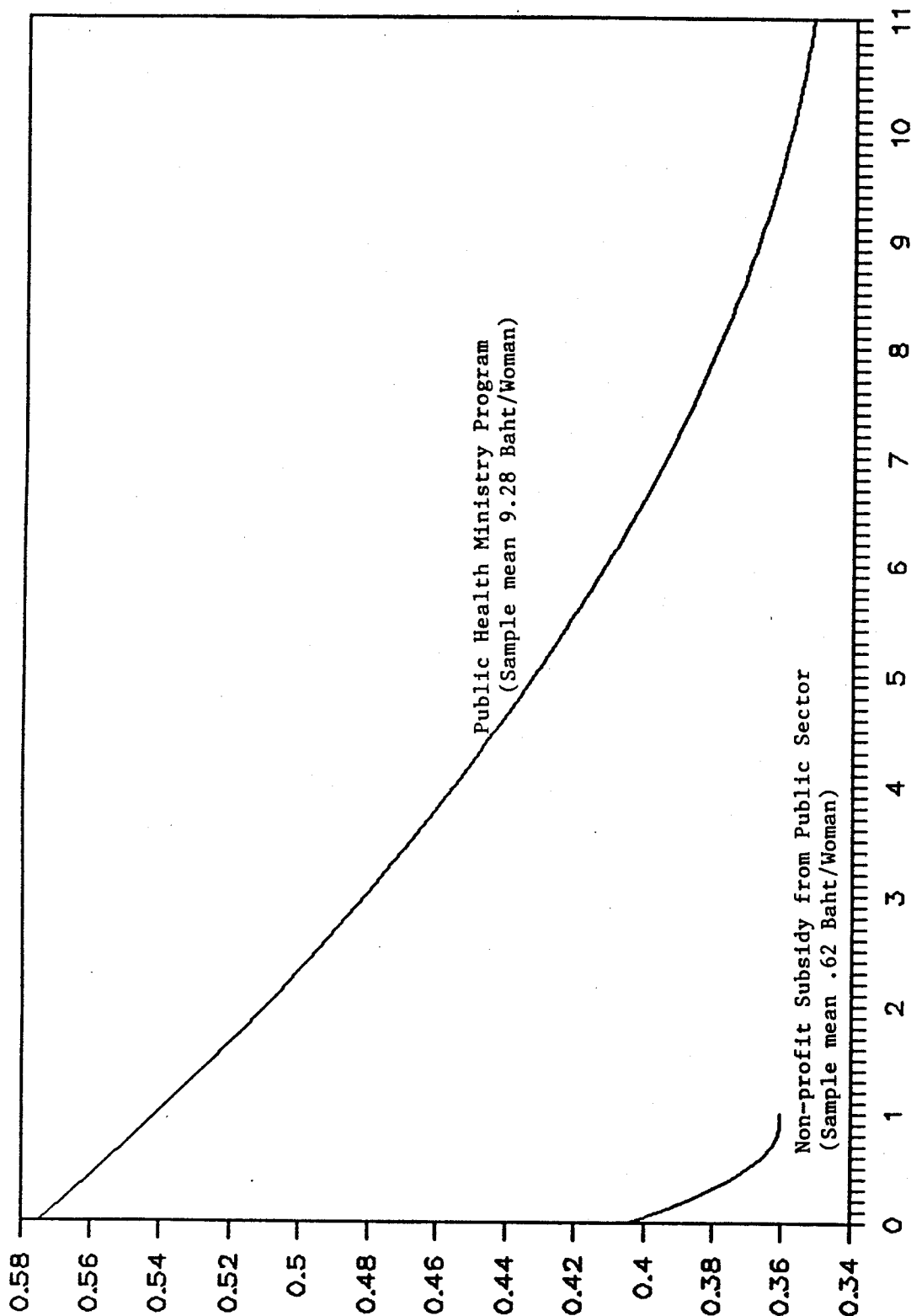
Because of the apparent positive effect on fertility of the interaction of private and public program expenditures, inputs to the private program are enhanced by a factor of three if they are concentrated in those areas that receive only half the national average level of public family planning outlays. 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 (3) 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 less than 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

FIGURE 3

LINEAR REGRESSION MODEL SIMULATING ASSOCIATION OF PROVINCE FAMILY PLANNING EXPENDITURES AND FERTILITY

Linear 0-4



PUBLIC EXPENDITURES IN BAHT PER WOMAN AGE 15 TO 49 IN 1975

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 SRS a

Explanatory Variables:	Poisson Model		Ordinary Least Squares		Sample Statistics ^b
	(1)	(2)	(3)	(4)	
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. See attached notes to Table 3.

in the 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 are 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, particularly those in age 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 as they influence fertility are estimated and 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 benefit disproportionately the less educated women but appear to depress the fertility most for the educated. This pattern is surprising unless the public sector activities are oriented to the better educated urban middle class and their growing demand in this period for sterilization. The permanent income (i.e. consumption) interactions are also jointly statistically significant at

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 (10.2)	-.0463 (.01)	-26.7 (1.78)
Age Woman (in years)	.959 (7.30)	-.0967 (0.42)	1.41 (2.10)
Age Squared ($\times 10^{-2}$)	-.185 (6.96)	.0277 (.06)	-1.78 (2.33)
Education woman (in years)	-.108 (16.9)	-.0233 (4.76)	-.0526 (2.26)
Household Monthly Consumption per adult (log) ^b	.726 (11.2)	.379 (4.77)	.302 (1.42)
Family Planning Activity (Bahts per woman in 1975)			
Public	-.0228 (3.26)	-.0154 (1.64)	-.0235 (1.13)
Private	-.152 (3.92)	-.120 (2.16)	-.0574 (.49)
Resident in Municipal Area	-.219 (4.89)	-.0873 (1.50)	-.497 (3.47)
Log likelihood	-4281.6	-2804.9	-1120.6
Sample Size	4490	3159	2485
Dependent Variable (standard deviation)	.604 (.757)	.482 (.682)	.163 (.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
² X Likelihood ratio statistical significance level over Col. (1) estimates Table 1	5.8	11.0	15.4	7.6	32.0
	.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.

the .01 percent level in col. 2 and 3, confirming that the lower income household gains more than the rich in terms of the number of births that the programs avert. 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 sector program than in the public sector.

The private program has a statistically significantly larger negative effect on fertility in nonmunicipal or rural areas than in the urban/municipal areas as seen from col. 4 of Table 4. The Northeast region of Thailand is the poorest; it has 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 income, rural residence, or living in the impoverished Northeast region of Thailand. Though statistically weak, these program interactions suggest a given budget for private non-profit public family planning expenditures would be more effective if it were 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 further 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. Because the fertility of less-education women, living in low income (or consumption) households, is most notably affected by

the family planning program, it would seem appropriate to devise a fee for service that price-discriminated in favor of these poorer segments of the population. With the interactions estimated in Table 4, it is possible to estimate how much of an effect could be expected on fertility by changing the emphasis of program from a nationally free system to one that was gratis only in certain regions of the country.

To explore this possibility, Table 5 reports the children 0-4 linear regressions separately for the urban, suburban and rural populations. Comparisons of effects confirm that the impact of women's education, permanent income, or family planning is not concentrated 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 been reduced in higher income urban populations, but the elasticity parameter is estimated to be twice as large in municipal than in rural or suburban Thai populations. But private non-profit family planning outlays are particularly effective outside of the municipal areas where the commercial market for contraceptives is less well established. The cross substitution of the private nonprofit and public family planning programs is also most notable in the rural villages where the commercial market for contraceptives is least well developed.

It is possible that omitted variables representing the community's education and health programs and other, unobserved constraints on reproduction that influence desired and actual fertility levels in the population may be correlated with the 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. There are 279 clusters in the sample, which are employed here and all variables are expressed as deviations from the community cluster mean of

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 ($\times 10^{-2}$)	-.182 (15.5)	-.201 (13.2)	-.254 (27.5)	-.222 (38.5)
Education Woman	-.0299 (11.7)	-.0248 (5.22)	-.0305 (8.21)	-.284 (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.

that variable. These estimates, as with the urban, suburban and rural strata, are very similar to the overall estimates based also on the between cluster variation in fertility (Table 1, Column 4). The F test of the statistical improvement gained by including the 279 regional effects is not statistically significant, confirming this source of omitted variable effects is not important in the linear regression 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 and 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, as confirmed in this study, their measure of "economic development" 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 total family income. But women's labor supply is determined simultaneously with fertility and is included in total family income or measured GNP per capita. Different sources of family income -- those derived from the value of the time of men and women and physical wealth -- embody different price and income effects on fertility, and thus exert different net effects. 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 associated with her

education and male productivity and sources of family nonearned income, including land ownership.

Another study of the allocation of family planning resources in Thailand found a large effect of program staff assignments on contraceptive use within the public program (ESCAP, 1987). 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 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 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. 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 are needed. The omission of intercorrelated control variables in aggregate analyses will bias the estimated partial effects of the remaining variables. Analysis of the individual level survey data permits one to estimate more flexible nonlinear relationships and test more detailed hypotheses that can clarify who benefits from public sector expenditures on family planning. The personal distribution of public program benefits is almost as important as the total program effects in evaluating their desirability.

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 in 1981, a country in which total fertility rates declined by nearly half in the two decades before the 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.^{14/} 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 is 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 planning program effects at the individual level, the own-children measure of women's recent fertility is an empirical compromise imposed by the source 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, to wait until lifetime cumulative fertility of cohorts can be linked to 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 (Chamrathirong, et al., 1987). Household income appears to relax the economic constraint on having

more children.^{15/} Although incomes are rising rapidly in Thailand, other factors making children more costly or less valued in large numbers have offset these gains in income. The increasing education of women is one development that has strongly countered the effect of increasing 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 that has 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 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 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.

Statistical Appendix

The 1981 SocioEconomic Survey sampled thirty amphoes (districts) in each of the four regions of Thailand outside of Bangkok, and thirty 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 a publicly released file.

A distinct feature of the 1981 SES is that it treated the sub-household with its own income as separate in so far 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 to children who might possibly have been their own children or to spouse in the household. Six instances 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 be normally maintained in separate households and then present fewer matching problems. Twenty-two men were children of heads, but had more than one female child in law in the household to whom they might have been married. Conversely, 36 female children of the head have multiple possible spouses related to the head as "children in laws." This attrition rate from the sample of 1.65 percent was smaller than I expected to encounter and is probably due to the emphasis given 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 half of the men reported these data that permit me to compute for them an hourly wage rate.

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

(Wage Participants/All Persons)

Notes:

^aThe 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

^aRegression model is specified as eq. (3) in Table 1.

Notes

^{1/} The complexity of dynamics and life cycle context in which fertility occurs suggests the desirability of measuring how public programs in family planning help couples have the children they want, when they want them, over their lifetime. In this paper, I examine the more limited question of how period fertility rates are affected by the program (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 standard Mincerian (1974) wage function specification. These calculations, however, do not take account of the possibly unrepresentative character of workers and wage earners. The estimates are then corrected for this dual 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 additionally 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. This rough approximation for the private rate of return to schooling (Mincer, 1974) for women in Thailand may require adjustment, if those who receive wages are not representative of the entire population. Following Heckman's (1979) approach, but allowing for two selection processes of entering the labor force and obtaining a wage earning job, adjusted estimates of the wage function for women still imply that an additional year of schooling is associated with a 18 percent increase in wages. However, these private returns to schooling for Thai women are not uniform across education levels. This selection-corrected measure of private returns

to schooling is lower than average at the primary level (at 10 percent) and higher than average at the secondary level (at 25 percent). The sample selection correction procedure has its largest effect on the estimates of returns to higher education, where the uncorrected return is 2 percent and the corrected return is 11 percent. If years of female education exerted their effect on fertility only by means of raising the value of women's wages, and hence the cost of child care, and assuming that this net effect of the woman's wage on fertility is proportional, then education at the secondary level would have a greater effect on fertility than either at the primary or university levels. To allow for this possibility, years of education is entered both linearly and in quadratic form in the fertility model.

^{3/} See Appendix Table A-2, excluding regional variables representing Bangkok municipal areas, sanitary districts and the Northeast region. When these regional shifters are included much of the private returns realized by education through rural-urban migration is eliminated from the estimate of education's proportional effect on wage rates.

^{4/} See Table A-2.

^{5/} Clearly, the woman's education may influence fertility for reasons other than its effect on the woman's market wage opportunities. Education could reduce the cost for a woman to evaluate and adopt new methods of birth control, change her pool of potential marriage partners, and modify her preferences for having children relative to other things, as well as influence her health and possibly fecundity. The current study does not discriminate among the alternative mechanisms that might explain the net negative relationship between the woman's years of schooling and her fertility. Recent studies of Thai fertility patterns do suggest, however, that educational differentials are decreasing in the 1970s and 1980s (Aphichat Chamrathirong, et al., 1987).

^{6/} See later discussion of the 1981 SES new concept of the nuclear family. See Carmel Chiswick, 1987.

^{7/} 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.

^{8/} 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.

^{9/} See Statistical Appendix text and specifically Table A-1.

^{10/} Others have developed regional estimates of fertility by analogous methods in Korea, China and Thailand. See for example the work of Arnold et al. 1985, based on the 1980 Census of Thailand.

^{11/} 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).

^{12/} Serial correlation over time in the unexplained error or propensity of some couples to have births 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.

^{13/} The likelihood function is

$$L = \sum_i \sum_t (c_{it}^{x_{it}\beta} - e^{x_{it}\beta} - \ln c_{it}!),$$

and the derivatives with respect to β are

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

^{14/} 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.

^{15/} There are few international comparisons of the frequency of child fosterage, except to note its high frequency in West Africa (Ainsworth, 1988). Child fosterage is not common in Thailand, and as indicated earlier the relevant child mortality rates are less than five percent if analysis deals with own children under the age of five in 1981 (Knodel, et al., 1987). Separation of mother and child to facilitate the child's school attendance should also be a relatively minor problem in Thailand, given the widespread availability of primary schools, until a child is more than 10 years of age.

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