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Working Paper

Socioeconomic Determinants of Fertility in Côte d'Ivoire

Center Discussion Paper, No. 557

Provided in Cooperation with:

Yale University, Economic Growth Center (EGC)

Suggested Citation: Ainsworth, Martha (1988) : Socioeconomic Determinants of Fertility in Côte d'Ivoire, Center Discussion Paper, No. 557, Yale University, Economic Growth Center, New Haven, CT

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CENTER DISCUSSION PAPER NO. 557

SOCIOECONOMIC DETERMINANTS OF FERTILITY IN COTE D'IVOIRE

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July 1988

Notes: This research was supported by the Living Standards Unit of the Development Research Department of the World Bank.

Center Discussion Papers are preliminary materials circulated to stimulate discussion and critical comments. References in publications to Discussion Papers should be cleared with the authors to protect the tentative character of these papers.

The opinions expressed are those of the author and do not reflect policy of the World Bank or its members. Comments on an earlier version from T. Paul Schultz, John Strauss, Duncan Thomas, and Vassilis Hajivassiliou of the Economic Growth Center, Yale University, and Dennis DeTray and John Newman of the World Bank are gratefully acknowledged.

ABSTRACT

This paper estimates a reduced form equation of the socioeconomic determinants of fertility in Côte d'Ivoire. The number of children ever born is regressed on the mother's age and schooling, the location of the household, and household income variables. This equation is estimated using ordinary least squares (OLS), maximum likelihood Tobit, and a Poisson count model. The advantages and drawbacks of the different econometric models in modelling fertility are discussed. Data are from 1444 women interviewed by the 1985 Côte d'Ivoire Living Standards Survey.

For the entire sample, female schooling has a depressing effect on fertility, while household income is associated with higher fertility. Among the subsample of urban women, only the negative effect of schooling is observed; among the subsample of rural women only the positive effect of household income is observed. The absence of a schooling effect among rural women is attributed in part to the low proportion of women with any schooling. When the sample is broken into three age cohorts, the negative effect of schooling on fertility is observed for the youngest and middle cohorts (ages 15-24 and 25-34, respectively), while the positive effect of income is observed for the middle and oldest cohorts (25-34 and 35+, respectively). This suggests that a fertility decline may be underway among young educated women.

The robustness of the results to the specification of income is also examined. Three income measures are used: the value of household consumption per adult (a proxy for permanent income); household income per adult; and household nonlabor income per adult. Results were most robust for the permanent income measure, less so for current income, and generally insignificant for nonlabor income.

I. Introduction

Subsaharan Africa is the poorest region of the world and the region with the highest birth rates. Whereas fertility is declining in every other developing region, there are no documented cases of national fertility decline in Subsaharan Africa, with the possible exception of Zimbabwe (World Bank 1986). Studies of the determinants of fertility in other developing regions have generally found that female schooling depresses fertility, with the effect of income more ambiguous (cf. Cochrane 1979, T.P. Schultz 1974, 1981). Studies of the economic determinants of African fertility have been limited by the lack of household data sets with adequate demographic and economic information.

This paper examines the likely impact of the spread of schooling and rising income on fertility in Côte d'Ivoire with data from the 1985 Côte d'Ivoire Living Standards Survey. Per capita income in Côte d'Ivoire was \$660 in 1985, making it a relatively prosperous country by African standards (World Bank 1987). GDP grew by an average 6.8 percent per year between 1965-80 but declined at a rate of 1.7 percent per year from 1980-85. The population of 10 million is growing at 3.8 percent per year (1980-85), the combined effect of a high rate of natural increase (3.1 percent annually) and immigration from neighboring countries. Roughly 30 percent of the population is foreign born and about 40 percent of the total population resides in urban areas. Schooling has spread rapidly since independence, but lags behind other Subsaharan countries with similar incomes. The Ivorian gross primary enrollment ratio for females is only 63 percent, compared to 94 percent for Kenya, 97 percent for Cameroon, and 127 percent for Zimbabwe.¹

Despite rapid economic growth, high urbanization, and the spread of schooling, the average woman in Côte d'Ivoire has 6-7 children by the end of her childbearing years. The 1980-81 Ivorian Fertility Survey found that only 3.8 percent of currently married, fecund women were using any method of family planning and only 0.6 percent were using an effective method such as the pill or IUD (R.C.I. 1984). The government has been pronatalist since independence and there are virtually no family planning services available except from private sources in the largest city, Abidjan.

This paper is organized as follows. Section II presents the economic model of fertility that motivates the reduced form equation. Section III describes the data set, the equation to be estimated, and construction of the income variables. Section IV discusses the three econometric models used in estimation and evaluates their appropriateness in analyzing fertility. Section V presents estimation results. The concluding section summarizes the findings.

II. A model of the determinants of fertility

The choice of variables in the reduced form equation is motivated by a theoretical model of the demand for children in the tradition of Becker (1965, 1981). The model adopts the perspective of an individual woman,² who maximizes a long-run, concave, twice-differentiable utility function over children (C), market goods (X), and her own leisure (L):

$$U = U (C , X , L), \quad U' > 0, U'' < 0 \quad (1)$$

The utility function is maximized subject to a household production function for children and to time and budget constraints. The production of children is described by a linearly homogeneous production function with

mother's time in childrearing (T_c) and purchased child goods (X_c) as inputs:

$$C = \gamma (T_c, X_c) \quad \gamma' > 0, \gamma'' < 0 \quad (2)$$

Since the model takes the perspective of individual women, married or not, the husband's time input does not enter this production function. This is a common assumption in modelling fertility in the U.S. and is probably more realistic in Subsaharan Africa.

The woman's time constraint allocates total time (Ω) among childrearing (T_c), market production (T_m), and leisure:

$$\Omega = T_c + T_m + L \quad (3)$$

The full-income budget constraint sets the total value of the woman's time plus nonlabor income equal to consumption "expenditure":

$$w\Omega + V = \pi_c C + p_x X + wL \quad (4)$$

where w = the woman's market wage; V = nonlabor income; π_c = the shadow price of children; and p_x = the price of other market goods. If we assume that the time allocation of other household members (such as the husband) is exogenously given and that their leisure does not enter the woman's utility, then their income can be considered exogenous and included in the woman's nonlabor income.

The shadow price of children is the sum of the value of their marginal inputs in production:

$$\pi_c = w t_c + p_{x_c} X_c \quad (5)$$

where t_c is the marginal time input in child production ($\delta T_c / \delta C$), p_{x_c} is the price of purchased child inputs, and x_c is the marginal input of purchased goods ($\delta X_c / \delta C$).

Maximizing equation (1) subject to the constraints (2) - (4), assuming that the first and second order conditions are met, yields equations expressing the demand for children, market goods, and leisure as a function of exogenous prices and nonlabor income. The comparative statics for the demand for children can be signed with some assumptions. If children and market goods are substitutes in consumption and the substitution effect exceeds the negative income effect of a price change, an increase in the price of market goods raises the demand for children. If we assume that children are normal goods, then an increase in nonlabor income will also raise the demand for children.⁵ The effect of an increase in the woman's wage on the demand for children can be divided into two components: the effect of an increase in the shadow price of children (which unambiguously lowers demand) and the effect of an increase in the price of leisure (which is negative if children and leisure are complements but ambiguous if they are substitutes). Empirical studies have generally found the net effect of the woman's wage on fertility to be negative, and that is what we expect here. The demand for children and posited signs are summarized as:

$$D_c = D_c (\overset{+}{p_x}, \overset{-}{p_{xc}}, \overset{-}{w}, \overset{+}{V}) \quad (6)$$

The number of children observed is the result of the interaction between the demand for and supply of children. The supply of children is biologically determined by the age of the woman (A) and a variable (μ) that measures a woman-specific component of fecundity (Rosenzweig and Schultz, 1985). The supply of children increases with age but at a decreasing rate, declining absolutely as the woman reaches the biological end of childbearing.

$$S_c = S_c (\overset{+}{A}, \overset{+}{\mu}) \quad (7)$$

The reduced form equation for the determinants of fertility includes both demand and supply-side variables, since there is insufficient information to separately identify supply and demand. Note that this long-run model assumes that women make a "once and for all" decision about the number of children to have, based on their perceived lifetime wage, income, and exogenous prices. Yet fertility decisions are clearly dynamic. Preferences change over the life cycle and expectations about the future may not be realized. With this cross-sectional data set, estimation of a dynamic model of fertility is not possible.

III. The data

Data are from the Côte d'Ivoire Living Standards Survey (CILSS), a permanent household survey begun in February 1985 by the Côte d'Ivoire Department of Statistics and the World Bank. The CILSS interviews 1600 households annually, spread out over 12 months. It obtains detailed socioeconomic data on all household members, complete consumption and income measures at the household level, and a fertility history from one randomly selected woman 15 years or older in each household. The survey methodology is documented in Ainsworth and Muñoz (1986) and Grootaert (1986).

Of the 1599 households surveyed between February 1985 and January 1986, 1488 had women age 15 or older. Twenty women were dropped from the analysis because of inconsistent fertility records and an additional 24 were dropped because either income or consumption variables had not been computed for their households.⁴ This leaves 1444 women for analysis.

The average woman in the sample is 34 years of age with a mean of 3.9 children.²⁵ Seventy-two percent of the women are currently married, 17 percent have never been married, and the remainder are divorced, widowed, or separated. Mean age at first cohabitation for the ever-married women in the sample is 17.6 years. Mean household size is 8.6 persons, of whom 4.6 are adults. The level of schooling in the sample is quite low -- only 24 percent of the women have had any schooling and mean schooling (including those with none) is a mere 1.7 years. The absence of any schooling is particularly severe for women over age 35 and for rural women, with only 4.5 and 10.5 percent, respectively, having had any schooling. Table 1 presents mean children ever born (CEB) by level of schooling, controlled for current age. The number of children increases with age and decreases with the amount of schooling.

TABLE 1: Mean children ever born by age and schooling

LEVEL OF SCHOOLING	AGE								TOTAL
	15-19	20-24	25-29	30-34	35-39	40-44	45-49	50+	
NONE	.74 (126)	1.92 (131)	3.67 (132)	4.70 (117)	5.90 (125)	6.08 (106)	6.50 (105)	6.02 (252)	4.53 (1094)
PRIMARY	.55 (69)	1.94 (52)	3.83 (35)	3.89 (28)	4.71 (7)	- (2)	- (2)	- (1)	2.22 (196)
SECONDARY AND HIGHER	.26 (46)	.84 (49)	1.90 (21)	3.05 (22)	4.73 (11)	- (4)	- (1)	.. (0)	1.66 (154)
TOTAL	.59 (241)	1.69 (232)	3.50 (188)	4.35 (167)	5.75 (143)	6.13 (112)	6.50 (108)	6.02 (253)	3.91 (1444)

Note: The number of observations is in parentheses. Means are not reported for cells with fewer than 5 observations.

The estimated reduced form equation regresses children ever born, the endogenous choice variable, on a set of five exogenous variables: age, age squared, years of schooling, urban residence, and one of three household income variables. The equation is estimated for the entire sample, for urban and rural women separately, and for three age cohorts.

Age and age squared are included to control for the biological supply of children. Since the reduced form is estimated for women of all ages, many of whom are still of childbearing age, these variables control for exposure to the risk of pregnancy.

Years of schooling is used as a proxy for female wages, which were not available for most women. As a proxy for wages, an increase in schooling should lower the demand for children. Maternal schooling may have independent effects on the demand for children other than as a proxy for wages, however. It may improve maternal health, raising the supply of children, or by improving child health (lowering child mortality) it may increase childspacing intervals and reduce the supply of children. If the mother's demand for children is really a demand for surviving children, the lower child mortality associated with mother's schooling may lead her to have fewer pregnancies. Schooling may also affect women's preferences, inducing them to demand fewer children of higher "quality".⁶

A dummy variable for urban residence is included to reflect greater wage-earning opportunities for women in urban areas.⁷ It also measures greater availability of services of all types -- market services, schools, health facilities, and other economic infrastructure. Urban residence should be associated with a higher shadow price of children and thus lower demand. Better health services in urban areas would lower the

supply of children through reduced child mortality and longer birth intervals, induced by the extended period of breastfeeding and postpartum amenorrhea if a child survives. Better urban health services could equally raise fecundity and the supply of children by better treatment of sexually transmitted diseases that would otherwise lead to infertility.

The theoretically correct income variable to use is "nonlabor income". It should exclude the woman's own earnings, which are endogenous through her labor supply. The earnings of other household members can be included, since their labor supply is considered exogenous to fertility decisions here. Unfortunately, except for the 5.5 percent of the women in the sample who had wage income, it was impossible to attribute income to individuals and to purge household income of the woman's earnings.

Three different household income variables are used in the empirical estimation. The principal income variable is a proxy for household permanent income that includes annual consumption expenditure, the value of home production consumed, and an imputed value of services from durables.²³ Consumption is used as a proxy for permanent income because it tends to fluctuate less over the life cycle than current income. The variable used is the natural log of "permanent income" as defined above, in CFA francs, per adult household member (15 years or older).

The two other household income variables are: (a) current income, the sum of income from wages, home agricultural production, home businesses, the value of services from durable goods, receipt of transfers, imputed rent for owners of housing in urban areas, and other income; and (b) nonlabor income, which includes the value of services from durable goods, receipt of transfers, income from rents on property, dividends, imputed rent for owners

of housing in urban areas, and all other household income not tied to labor supply. Social security and pension income are not included in household nonlabor income. Both income variables have been divided by the number of adults in the household and are expressed as natural logarithms.²⁹

All three of the income variables have major shortcomings. The permanent and current income variables suffer from endogeneity, since the woman's consumption and earnings could not be netted out from the rest of the household. Further, although dividing through by the number of adults makes the permanent income variable less dependent on the left hand side variable, children ever born, the presence of children in the household will nevertheless drive up the level of consumption per adult. Ten percent of all households and 15 percent of rural households had no nonlabor income, and the value of owner-occupied housing could be assessed only for urban households. The nonlabor income variable is also sensitive to the assumptions used to value the services from housing and durable goods.

Landholdings are not included as a regressor because ownership is not well defined in much of land-abundant rural Côte d'Ivoire. The land cultivated is not an acceptable alternative because it is endogenous. Respondents were generally unable to cite with accuracy the area owned or cultivated and in most of the rural communities there is no land market, making valuation of land impossible.

Exogenous health service variables, such as the distance to the nearest maternity ward, maternal and child health clinic, and hospital, were tentatively included in the regressions but subsequently dropped due to problems in interpreting their coefficients. In rural areas, services of all types tend to be clustered at the same administrative level; the

distance to a health facility is also the distance to all economic infrastructure and services. Access to services on the date of the interview is unlikely to be a good proxy for lifetime access to health services, which is clearly more relevant to fertility decisions. Finally, access to health services may be endogenous if people selectively migrate to areas with better service availability. Alternatively, the government may place health services in areas with the worst health and highest fertility. For all of these reasons, it is difficult to interpret coefficients on distances as the effect of access to health services.¹⁰

The means and standard deviations of the dependent and independent variables are presented in Table 2.

IV. Estimation techniques

The reduced form equation is estimated using three econometric models: ordinary least squares; maximum likelihood Tobit; and a Poisson count model.

Ordinary Least Squares. Least squares estimates have been widely used in fertility analysis (see Anker and Knowles 1982 and the studies cited in Cochrane 1979, T.P. Schultz 1974, 1981, and T.W. Schultz 1974). OLS expresses the dependent, continuous random variable (y) as a linear function of exogenous variables (x) plus an error term (ϵ), where it is assumed that ϵ is independent, identically distributed, and uncorrelated with the regressors.

TABLE 2: Sample means and standard deviations

VARIABLE	All women		Urban		Rural		Age 15-24		Age 25-34		Age 35 +	
	Mean	SD	Mean	SD	Mean	SD	Mean	SD	Mean	SD	Mean	SD
Children ever born	3.91	3.30	3.14	3.05	4.46	3.37	1.13	1.28	3.90	2.27	6.06	3.29
Age	34.31	15.07	30.38	12.93	37.07	15.85	19.53	2.73	28.94	2.73	48.74	11.38
Age squared	1403.8	1266.7	1090.2	1019.0	1624.9	1373.6	388.78	106.70	844.84	159.00	2505.3	1246.4
Years of schooling	1.69	3.43	3.40	4.47	0.48	1.59	3.06	3.88	2.26	4.20	0.31	1.56
Urban dummy	0.41	0.49	1.00	0.00	0.00	0.00	0.53	0.50	0.45	0.50	0.30	0.46
Ln permanent income/adult	12.59	0.82	13.09	0.72	12.24	0.68	12.65	0.77	12.77	0.88	12.44	0.79
Ln current income/adult	12.29	1.43	12.81	1.33	11.92	1.39	12.44	1.00	12.47	1.52	12.04	1.62
Ln nonlabor income/adult	8.31	3.32	10.10	2.52	7.04	3.23	8.76	2.99	8.42	3.17	7.89	3.59
N	1444		597		847		473		355		616	

Children ever born is not a continuous variable, however, and it is censored at zero. When the dependent variable is censored, least squares estimates are inconsistent because the error term is not independent of the regressors (Amemiya 1984, Maddala 1983). When both y and the regressors are normally distributed, OLS coefficients are biased downward in proportion to the proportion of nonzero observations (Amemiya 1984, Greene 1981, Maddala 1983). Thus, the smaller the degree of censoring, the closer are OLS and Tobit coefficients. All of the OLS specifications exhibited heteroskedasticity and the standard errors were re-estimated using the White heteroskedastic-consistent covariance matrix.

Maximum Likelihood Tobit. Let y^* be the true demand for children (which can be positive or negative) and y the observed number of children ever born, which can take a value of zero or positive integers. The Tobit model takes into account the fact that negative values of y^* are not observed, but it still assumes that y is a continuous variable in the nonnegative range. The model is:

$$y^* = x'\beta + \epsilon$$

$$y = y^* \text{ if } y^* \geq 0 \\ = 0 \text{ otherwise}$$

$$\epsilon \sim N(0, \sigma^2)$$

$$E(\epsilon) = 0 ; E(x'\epsilon) = 0$$

$$E(\epsilon_i \epsilon_j) = \sigma^2 I, \quad i = j \\ = 0, \quad i \text{ not equal } j$$

$$E(y^*) = x\beta$$

Tobit coefficients are estimated using maximum likelihood methods.¹¹ The Tobit ML estimator is the most efficient consistent and asymptotically normal estimator but is sensitive to the assumptions of homoskedasticity and normality of the errors (Maddala 1983, Amemiya 1984). Violation of either

of these assumptions produces inconsistent estimates, and the direction of the bias is ambiguous.

Poisson count model. The Poisson count model assumes that the dependent variable is generated from a Poisson process and takes on values that are nonnegative integers. Thus, both the censored and integer aspects of children are taken into account. The Poisson model has the additional advantage that it models some heteroskedasticity, since the variance of the dependent variable is a function of $x'\beta$. The model is:

$$\text{Pr } (y_t) = \frac{\exp(-\lambda_t) \lambda_t^Y}{Y_t !}$$

where $\lambda_t = \exp(x'\beta) = E(y_t) = \text{variance}(y_t)$. The Poisson model is estimated using maximum likelihood methods.¹²

One problem with the Poisson model is that it restricts the mean of the dependent variable to equal the variance. Violation of this restriction produces overly small standard errors on the coefficients (Portney and Mullahy 1986, Cameron and Trivedi 1986). For the data set used here, the variance/mean ratio for children ever born is 2.79 to 1 (although for some subsamples this ratio is less than 2 to 1). There are two approaches to the problem of "overdispersion" of the data. The first is to model the overdispersion. Hausman, Hall, and Griliches (1981) use a negative binomial model in their study of patents and research and development expenditure, for example. Cameron and Trivedi (1986) compare results from Poisson and various negative binomial models, however, and find that even in the presence of overdispersion the Poisson and negative binomial coefficients are quite similar, although the Poisson model yields overly-small standard errors.

The second approach to the overdispersion problem is to use a method for estimating the standard errors on the Poisson covariance matrix of β that is more robust to violation of the restriction that the mean equals the variance. This is the approach adopted here; the standard errors reported for Poisson regressions are the "robust" standard errors described in Portney and Mullahy (1986).¹³

V. Estimation results

Estimates of the reduced form equation for all women and various subsamples are presented in Tables 3 and 4. The first figure in each cell is the coefficient (β) for the variable in the regression. To facilitate comparisons across models, the slope coefficients on the expected value functions of the Tobit and Poisson models have been calculated at the sample means and are presented in brackets; it is these figures that should be compared with the OLS coefficients.¹⁴

Reduced form estimates for the entire sample using the permanent income variable as a regressor are in the first three columns of Table 3. All of the coefficients are highly significant and of the expected signs. At the mean, an additional year of schooling reduces the number of children ever born by about 0.14. Experimentation with other specifications of schooling revealed that the effect of schooling is nonlinear, concave, and negative for all positive years of schooling, including the first year of primary school.¹⁵ Thus, even low levels of primary schooling have a negative effect on fertility in Côte d'Ivoire. The coefficients on

TABLE 3: Estimates by location

Dependent variable: Children ever born

EXPLANATORY VARIABLES	ALL WOMEN			URBAN WOMEN			RURAL WOMEN		
	OLS	Tobit	Poisson	OLS	Tobit	Poisson	OLS	Tobit	Poisson
Age	0.4296** (.0218)	0.5414** (.0240) [0.4842]	0.1379** (.0077) [0.4286]	0.4891** (.0308)	0.6686** (.0362) [0.5420]	0.2009** (.0301) [0.4815]	0.4106** (.0285)	0.4106** (.0280) [0.3780]	0.1163** (.0084) [0.4862]
Age ²	-0.0038** (.0003)	-0.0050** (.0003) [-0.0045]	-0.0013** (.0001) [-0.0040]	-0.0046** (.0004)	-0.0066** (.0004) [-0.0053]	-0.0020** (.0004) [-0.0048]	-0.0036** (.0003)	-0.0036** (.0003) [-0.0033]	-0.0010** (.0001) [-0.0042]
Years of schooling	-0.1113** (.0158)	-0.1562** (.0356) [-0.1397]	-0.0443** (.0071) [-0.1377]	-0.0990** (.0168)	-0.1428** (.0355) [-0.1158]	-0.0366** (.0141) [-0.0877]	-0.0633~ (.0391)	-0.0633 (.0781) [-0.0583]	-0.0238 (.0169) [-0.0996]
Urban dummy	-0.4467** (.1606)	-0.6259** (.1881) [-0.5598]	-0.1338** (.0425) [-0.4158]						
Ln perm. inc/adult	0.3215** (.1015)	0.4287** (.1095) [0.3834]	0.1091** (.0260) [0.3391]	-0.0842 (.1286)	0.0391 (.1622) [0.0317]	0.0158 (.0745) [0.0379]	0.5827** (.1444)	0.5827** (.1242) [0.5365]	0.1399** (.0315) [0.5848]
Constant	-9.0986 (1.233)	-12.9165 (1.414)	-3.015 (.3415)	-5.2096 (1.588)	-10.5349 (2.164)	-3.132 (.9922)	-12.0359 (1.810)	-12.0359 (1.626)	-2.956 (.4230)
R ²	.44			.53			.37		
Sigma		2.7961 (.0509)			2.4879 (.0725)			2.6903 (.0541)	
LogL		-3119.8	-3181.2		-1143.5	-1144.5		-1969.3	-2005.4
N	1444	1444	1444	597	597	597	847	847	847

- Notes: 1. The first figure in every cell is the coefficient of the model. Standard errors are in parentheses. Figures in brackets are slope coefficients of the expected value functions, calculated at the mean (see footnote 13).
2. OLS standard errors are corrected for heteroskedasticity using the White heteroskedastic-consistent covariance matrix; Poisson standard errors are corrected to account for overdispersion (see text).
3. ** significant at .01; * significant at .05; ~ significant at .10.

permanent income per adult range from 0.32 to 0.38 and correspond to income per adult elasticities of +0.082 to +0.098.

The estimates for urban and rural subsamples reveal that the schooling coefficient is significant only for urban women (ranging from -0.09 to -0.12), while the income coefficient is significant only for rural women. The permanent income per adult elasticity for rural women ranges from +0.12 and +0.13 at the mean. That is, a ten percent increase in income per adult at the mean is associated with an increase in the number of children by about one and a quarter percent. The weak results for schooling of rural women are probably due to the very low levels of schooling and consequent low variation in the subsample. Alternatively, schooling may not be a good proxy for wages in rural areas where there are few wage-earning opportunities for women. An F-test on the OLS regressions found significant structural differences at the .01 level between the coefficients for urban and rural subsamples: $F(5,1434) = 4.75$. Likelihood ratio tests on Tobit and Poisson estimates confirmed this conclusion (LRT = 96 and 79, respectively).

The results in Table 3 show that schooling has a negative effect on current fertility but not necessarily on completed fertility. Women at different points in the life cycle are included in the regressions; younger women may be marrying later but having just as many children as their mothers by reducing childspacing intervals. Estimates by age cohort are presented in Table 4.¹⁶

Among the oldest cohort (35+ years), schooling has had no discernable effect on fertility but the effect of income is positive and

TABLE 4: Estimates by age group

Dependent variable: Children ever born

EXPLANATORY VARIABLES	AGE 15-24			AGE 25-34			AGE 35 +		
	OLS	Tobit	Poisson	OLS	Tobit	Poisson	OLS	Tobit	Poisson
Age	-0.0226 (.2690)	1.0834* (.4875) [0.6526]	1.1160** (.3030) [0.9709]	0.4340 (.8734)	0.4340 (.9422) [0.4194]	0.1960 (.2184) [0.7473]	0.1950* (.0899)	0.1950* (.0854) [0.1883]	0.0335* (.0161) [0.2048]
Age ²	0.0066 (.0071)	-0.0177 (.0123) [-0.0107]	-0.0220** (.0074) [-0.0191]	-0.0038 (.0151)	-0.0038 (.0161) [-0.0037]	-0.0024 (.0037) [-0.0092]	-0.0018* (.0008)	-0.0018* (.0008) [-0.0017]	-0.0003* (.0002) [-.0018]
Years of schooling	-0.0685** (.0130)	-0.1242** (.0245) [-0.0748]	-0.0666** (.0140) [-0.0579]	-0.1565** (.0268)	-0.1565** (.0384) [-0.1512]	-0.0466** (.0088) [-0.1777]	-0.0700 (.0770)	-0.0700 (.1064) [-0.0676]	-0.0121 (.0140) [-0.0740]
Urban dummy	-0.3359** (.1177)	-0.5312** (.1920) [-0.3201]	-0.2952** (.1037) [-0.2568]	-0.2990 (.2766)	-0.2990 (.2828) [-0.2890]	-0.0744 (.0686) [-0.2837]	-0.6354~ (.3340)	-0.6354* (.3099) [-0.6135]	-0.1056~ (.0560) [-0.6455]
Ln perm. inc/adult	0.0719 (.0742)	0.1003 (.1251) [0.0604]	0.0712 (.0676) [0.0619]	0.3245* (.1678)	0.3245~ (.1695) [0.3136]	0.0858* (.0434) [0.3271]	0.4253* (.2020)	0.4253* (.1807) [0.4107]	0.0703* (.0334) [0.4297]
Constant	-1.5207 (2.651)	-14.3497 (5.062)	-13.92 (3.082)	-9.1276 (12.79)	-9.1276 (13.08)	-3.263 (3.259)	-4.0964 (3.531)	-4.0965 (3.326)	0.0905 (.6024)
R ²	.33			.14			.02		
Sigma		1.5485 (.0728)		2.1214 (.0645)			3.2758 (.0958)		
LogL		-639.4	-587.4		-766.0	-760.4		-1590.3	-1695.4
N	473	473	473	355	355	355	616	616	616

- Notes: 1. The first figure in every cell is the coefficient of the model. Standard errors are in parentheses. Figures in brackets are slope coefficients of the expected value functions, calculated at the mean (see footnote 13).
2. OLS standard errors are corrected for heteroskedasticity using the White heteroskedastic-consistent covariance matrix. Poisson standard errors are corrected to account for overdispersion (see text).
3. ** significant at .01; * significant at .05; ~ significant at .10. Age and age squared are jointly significant at .01 for all models and all subsamples.

significant. The absence of a schooling effect is not surprising, given the very limited amount of schooling in this cohort (mean of 0.3 years). The Tobit and OLS coefficients are virtually identical, as only 5 percent of the women have never had a live birth. Expected value coefficients are remarkably similar across models.

Among the youngest cohort (15-24), only the schooling and urban dummy coefficients are significant. The negative relation between schooling and fertility is probably due to the effect of schooling on delayed marriage. Many of these women are still in school or not yet married, and thus have not begun childbearing. They may also be delaying marriage or childbearing because of the income-earning opportunities in urban areas. The OLS and Tobit coefficients differ greatly due to the high degree of censoring (42 percent of this cohort have had no children). The expected value coefficients show greater spread across models for this cohort, the Poisson estimates being the smallest absolute value.

Most of the women in the middle cohort (25-34) are married and have borne children -- almost 4, on average. They have been married for enough time to have compensated for any delay in marriage by having children at closer intervals. The schooling coefficient for the middle cohort is negative, highly significant, and 2-3 times the size of the coefficient for younger women in absolute value. This finding is consistent with the hypothesis that a fertility decline is underway among educated women. The permanent income coefficient for women 25-34 is also positive and significant and corresponds to a larger income elasticity (+0.08) than for the oldest cohort. The urban dummy variable is not significant as it is for the youngest and oldest groups. Poisson coefficients on schooling and

income are greater in absolute value than Tobit and OLS, but the differences are not great.

Table 5 summarizes the point elasticities of fertility with respect to schooling and income for all subsamples in Tables 3 and 4. It is difficult to evaluate these elasticities, as potentially comparable studies include different sets of regressors, different measures of dependent and independent variables, and are often based on aggregated rather than individual data. The female schooling elasticity of approximately -0.06 falls at the low end (in absolute value) of schooling elasticities for developing countries surveyed by T.P. Schultz (1974). Note that schooling elasticities for urban women and for the youngest cohort (15-24) are two to three times as large in absolute value as those for the entire sample, however. The schooling elasticity for the entire sample implies that raising mean schooling from 1.7 to 3 years, holding all other variables constant, would lower mean fertility from 3.91 to about 3.73 children. The income elasticity of about $+0.09$ implies that a 10 percent increase in per adult permanent income would raise mean fertility from 3.91 to 3.95 children.

TABLE 5: Point elasticities

SUBSAMPLE	Permanent income			Schooling		
	OLS	Tobit	Poisson	OLS	Tobit	Poisson
All women	.082**	.098**	.087**	-.048**	-.060**	-.059**
Urban				-.107**	-.125**	-.095**
Rural	.131**	.120**	.131**	-.007~		
Age 15-24				-.185**	-.202**	-.156**
Age 25-34	.083*	.080~	.084*	-.091**	-.088**	-.103**
Age 35+	.070*	.068*	.071*			

Note: Elasticities calculated at the mean using OLS coefficients and Tobit and Poisson expected value function coefficients in brackets in previous tables. Elasticities are not reported for insignificant coefficients. ** indicates significant at .01; * indicates significant at .05; ~ indicates significant at .10.

The sensitivity of results to the choice of income variable is presented in Table 6. The coefficients on current income are positive, but about a third the size of the coefficients on permanent income. Only the Tobit current income coefficient attains reasonable significance. The fact that current income has a positive, if weaker, effect compared to permanent income is reassurance that the coefficient on permanent income is not simply reflecting the positive correlation between per adult consumption and the number of children in the household. The stronger effect of the permanent income measure indicates that fertility is more sensitive to permanent than current income, as theory would suggest. It might also reflect endogeneity of the permanent income variable. None of the coefficients on nonlabor income are significant.¹⁷

TABLE 6: Sensitivity of results to income specification

Dependent variable: Children ever born

EXPLANATORY VARIABLES	Permanent Income			Current Income			Nonlabor Income		
	OLS	Tobit	Poisson	OLS	Tobit	Poisson	OLS	Tobit	Poisson
Age	0.4296** (.0218)	0.5414** (.0240) [0.4842]	0.1379** (.0077) [0.4286]	0.4362** (.0217)	0.5497** (.0240) [0.4974]	0.1393** (.0240) [0.4455]	0.4375** (.0217)	0.5512** (.0241) [0.4988]	0.1395** (.0078) [0.4499]
Age ²	-0.0038** (.0003)	-0.0050** (.0003) [-0.0045]	-0.0013** (.0001) [-0.0040]	-0.0039** (.0003)	-0.0051** (.0003) [-0.0046]	-0.0013** (.0002) [-0.0042]	-0.0039** (.0003)	-0.0051** (.0003) [-0.0046]	-0.0013** (.0001) [-0.0042]
Years of schooling	-0.1113** (.0158)	-0.1562** (.0356) [-0.1397]	-0.0443** (.0071) [-0.1377]	-0.1001** (.0155)	-0.1385** (.0352) [-0.1253]	-0.0397~ (.0203) [-0.1270]	-0.0966** (.0152)	-0.1342** (.0352) [-0.1214]	-0.0383** (.0070) [-0.1245]
Urban dummy	-0.4467** (.1606)	-0.6259** (.1881) [-0.5598]	-0.1338** (.0425) [-0.4158]	-0.2882* (.1473)	-0.3999* (.1783) [-0.3619]	-0.0762* (.0504) [-0.2437]	-0.3080* (.1518)	-0.4196* (.1919) [-0.3797]	-0.0799* (.0405) [-0.2577]
Ln income per adult	0.3215** (.1015)	0.4287** (.1095) [0.3834]	0.1091** (.0260) [0.3391]	0.0965~ (.0566)	0.1088* (.0519) [0.0985]	0.0295 (.0187) [0.0944]	0.0305 (.0243)	0.0336 (.0246) [0.0304]	0.0076 (.0058) [0.0245]
Constant	-9.0986 (1.233)	-12.9165 (1.414)	-3.015 (.3415)	-6.4398 (.7365)	-9.1244 (.8258)	-2.056 (.6460)	-5.5227 (.3952)	-8.0849 (.5606)	-1.756 (.1565)
R ²	.44			.44			.44		
Sigma		2.7961 (.0509)			2.8054 (.0511)			2.8077 (.0512)	
LogL		-3119.8	-3181.2		-3125.2	-3191.9		-3126.2	-3195.1
N	1444	1444	1444	1444	1444	1444	1444	1444	1444

- Notes: 1. The first figure in every cell is the coefficient of the model. Standard errors are in parentheses. Figures in brackets are slope coefficients of the expected value functions, calculated at the mean (see footnote 13).
2. OLS standard errors are corrected for heteroskedasticity using the White heteroskedastic-consistent covariance matrix. Poisson standard errors are corrected to account for overdispersion (see text).
3. ** significant at .01; * significant at .05; ~ significant at .10.

VI. Conclusions

In a cross-section of Ivorian women of all ages, female schooling is associated with lower fertility while household income acts in the opposite direction. Raising women's schooling from a mean of 1.7 to 3.0 years (an increase of 76 percent) would lower mean children from 3.9 to 3.7. With such low levels of schooling in the sample, it is risky if not impossible to estimate what impact universal female primary schooling -- a quadrupling of the mean -- would have on fertility. Nevertheless, additional schooling was found to have a negative effect on fertility for all levels of schooling, even the early years of primary school. The fact that schooling coefficients for the youngest women are negative and highly significant and remain so into the 25-34 year age group is consistent with the hypothesis that a fertility decline may be underway among women with schooling.

Experimentation with different income measures found that the consumption-based proxy for permanent income has a stronger relation with fertility than does current income and that nonlabor income has no effect. Measurement of nonlabor income is far from ideal, however, having been based primarily on the valuation of services from durables.

Finally, experimentation with three econometric models revealed that OLS, Tobit, and Poisson expected value coefficient estimates were similar in magnitude and significance, despite the fact that OLS estimates are inconsistent if the data are censored and Tobit estimates are sensitive to heteroskedasticity or nonnormality of the errors.

Footnotes

1. The gross primary enrollment ratio for females is the total number of females attending primary school as a percentage of all females of primary school age. Since older children are often enrolled in lower classes, this ratio may exceed 100 percent.
2. Treatment of the woman as the decision-making unit is more than a convenient assumption in the African context. Recent work by Oppong and Bleek (1982) and Etienne (1979) on matrilineal societies in Ghana and Côte d'Ivoire, respectively, underline the self-sufficiency of women and the "marginality of men". Children automatically belong to the mother's lineage and extramarital fertility is not discouraged in the urban area of southern Ghana studied by Oppong and Bleek. Etienne finds that Baulé women behave as "autonomous social agents" and may even adopt the children of relatives as their own dependents, not to be shared with their husbands.
3. Becker (1981), p. 102, points out that if there is an interaction between the quality and quantity of children, the effective price of children might increase with income and the sign on income could be negative.
4. The latter omissions were random; income and consumption data for these households were inadvertently left off the original data tape. They have since been added, but the variables have not yet been computed.
5. Summary statistics for the sample are not nationally representative because of the selection process for the women: they were randomly selected, one per household. Women from larger households thus had a smaller probability of being included. The results must be weighted to get nationally representative figures. The nonrepresentativeness of the sample does not invalidate conclusions on relationships between fertility and socioeconomic variables.
6. The tradeoff between the quantity and quality of children is not studied here.
7. The urban dummy is treated as exogenous, although the household conceivably could have moved to an urban area in order to take advantage of work opportunities and greater availability of services. This would introduce a self-selection problem into the estimation, where women who prefer working in the market and having fewer children move to urban areas and those who prefer home production and more children stay in rural areas.
8. The household income and consumption variables were computed by Kozel. Details are in her 1987 dissertation.
9. All households had positive permanent income; 9 had zero or negative current income and 144 had no nonlabor income. Per adult income values less than or equal to zero were arbitrarily assigned a value of 1 CFA franc (about half a cent) before conversion into logarithms.

10. The distance to health and schooling services was available for rural households only. The results for these variables, when added to the variables of the reduced form for 773 rural women, are as follows:

$$0.0479 \text{ DPRIM} + 0.0082 \text{ DSEC} + 0.0100 \text{ DHOSP} + 0.0036 \text{ DPMI} - 0.0197 \text{ DMAT}$$

$$(.0577) \quad (.0073) \quad (.0067) \quad (.0026) \quad (.0080)$$

where DPRIM is the distance to the nearest primary school, DSEC the distance to the nearest secondary school, DHOSP the distance to the nearest hospital, DPMI the distance to the nearest maternal and child health clinic, and DMAT the distance to the nearest maternity ward, all in kilometers. Standard errors corrected for heteroskedasticity are in parentheses.

11. The Tobit log likelihood function is:

$$\text{Log L} = d \cdot \ln \left[(1/\sigma) \phi \left(\frac{y - x'\beta}{\sigma} \right) \right] + (1 - d) \ln \left[1 - \Phi \left(\frac{x'\beta}{\sigma} \right) \right]$$

where $d = 1$ when $y^* \geq 0$ and $d = 0$ when $y^* < 0$, ϕ and Φ are the normal density and distribution functions, respectively, and the t subscript denoting the observation has been left off of d , y , and x for convenience.

12. The Poisson log likelihood function, first (g) and second (h) derivatives are:

$$\begin{aligned} \text{Log L} &= - \sum \exp(x'\beta) + \sum y \cdot (x'\beta) - \sum \log(y!) \\ g &= \sum x \cdot (y - \exp(x'\beta)) \\ h &= - \sum \exp(x'\beta) \cdot xx' \end{aligned}$$

where the t subscript denoting the observation has been left off. The first order conditions are set equal to zero and solved iteratively using the Newton-Raphson method. The likelihood function is globally concave, so convergence to a global maximum is relatively rapid.

13. The coefficients of the Tobit model represent $\delta y^*/\delta x$. The coefficients for $E(y)$ are obtained as follows (Maddala 1983, Rosenzweig and Schultz 1985):

$$\begin{aligned} E(y) &= \text{Pr}(y > 0) \cdot E(y|y > 0) + \text{Pr}(y = 0) \cdot E(y|y = 0) \\ &= \Phi(\beta'x + \sigma \phi/\Phi) + (1 - \Phi) \cdot 0 \\ &= \Phi \beta'x + \sigma \phi \end{aligned}$$

$$\delta E(y)/\delta x_i = \left[\Phi + \frac{\beta'x}{\sigma} \phi - \frac{\beta'x}{\sigma} \phi \right] \beta_i$$

$$\delta E(y)/\delta x_i = \Phi \beta_i$$

where ϕ and Φ are the normal density and distribution functions of $\beta'x/\sigma$, evaluated at the mean, and the i subscript denotes an explanatory variable. The coefficients for the expected value locus of the Poisson model are calculated as follows:

$$E(y) = \exp(\beta'x)$$

$$\delta E(y)/\delta x_i = \beta_i \cdot \exp(\beta'x)$$

where $\exp(\beta'x)$ is calculated at the mean.

14. Portney and Mullahy (1986, p. 37) obtain consistent estimates of the covariance of estimated β using:

$$I(\beta)^{-1} \left[\sum_t (\delta l_t / \delta \beta) (\delta l_t / \delta \beta)' \right] I(\beta)^{-1}$$

where $-\delta^2 l / \delta \beta \delta \beta'$ is denoted as $I(\beta)$, l_t is the contribution of the "t"th observation to the log-likelihood function and the expression is evaluated at maximum likelihood estimates of β .

15. A quadratic specification of schooling revealed that fertility "peaks" at 0.4 years of schooling and declines thereafter. Numerous specifications with dummy variables for individual years of schooling were also examined, with the following OLS results:

(1)	- 0.2505~DUM16	- 1.0866***DUM7PL			
	(.1366)	(.1839)			
(2)	- 0.4461*DUM12	- 0.1940 DUM36	- 1.0813***DUM7PL		
	(.2069)	(.1547)	(.1843)		
(3)	- 0.6352*DUM1	- 0.2370 DUM2	+ 0.3472 DUM3	- 0.1013 DUM4	
	(.2484)	(.3128)	(.5566)	(.3581)	
	+ 0.1536 DUM5	- 0.3418~DUM6	- 1.0807***DUM7PL		
	(.2965)	(.1790)	(.1842)		

where: DUMn = dummy variable for n years of schooling; DUMmn = dummy variable for m through n years of schooling; DUM7PL = dummy variable for 7 or more years of schooling; the left-out variable is no schooling; and other variables in the regressions included age, age squared, dummy for urban residence, and the logarithm of permanent income per adult. OLS standard errors are corrected for heteroskedasticity and are in parentheses. A comparable study of the determinants of fertility in Kenya, which also controlled for income, found that primary schooling had no effect on fertility, while all schooling above primary level had a significant negative effect (Anker and Knowles 1982). That study, based on a 1974 household survey, was confined to currently married women, while the Côte d'Ivoire results reported here are for all women regardless of marital status. Further, the proportion of Kenyan women with any schooling in 1974 (45 percent) was almost double the proportion for Côte d'Ivoire in 1985 (24 percent).

16. Tests for structural differences in regressions for age subsamples were significant at the .05 level for OLS coefficients ($F(12,1426) = 2.07$) and .01 for Tobit and Poisson coefficients (LRT statistics of 248.2 and 276, respectively).

17. A more relevant question for many fertility data sets with no income variables at all is the effect on the schooling coefficient when income cannot be controlled for. Excluding income from the regression for all women reduces the schooling coefficient by about .02 in absolute value, compared with the permanent income specification. When income is excluded from the regression for rural women, the schooling coefficient loses significance altogether, leading to the erroneous conclusion that schooling has no effect on fertility.

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