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Abstract. A significant positive association is established between fertility and purchased calories per adult equivalent using individual household data for Kinshasa, Zaïre. The analysis is based on a sample of women aged 20-24, none of whose children have died. Characteristics likely to influence fertility both behaviorally and biologically, such as education, are included as separate independent variables. Both closed birth intervals and conditional birth probabilities are analyzed, with the latter providing more significant results.

Nutrition has been hypothesized to influence fertility in a number of ways. Until recently, the most commonly examined relationship pertains to the generally accepted positive association between infant mortality and fertility. Prolonged malnutrition of the mother has been shown to increase infant mortality by reducing the birth weight of children [Burke (1945) and Habicht (1973)]. Improvements in nutrition, therefore, will lead to declines in fertility through decreases in infant mortality.

There are at least two other means by which choices involving nutrition may be expected to affect fertility. One is behavioral; the other is biological. In most low-income countries, it is argued that birth control is practiced only during the period immediately after pregnancy through the adoption of intercourse taboos during lactation [Ascadi (1974) and Tabbarah (1971)]. The purpose of birth control in this context is to prevent a new conception during the desired period of lactation, in order to increase the survival chances of the child who is already alive. The effect is to reduce the total cost of nutriton for the household, particularly in a situation where the relative price of human milk substitutes is high. The lower the income and hence nutrition level of the household, the more its satisfaction can be increased by lengthening the lactation period in this way. From this argument it follows that, with relative food prices and infant mortality rates constant, real income growth and associated nutrition increases may be expected to have a pro-natalist effect through a reduced incentive to abstain from intercourse during lactation.

The other mechanism involves the biological relation of nutrition to fertility and also suggests the possibility of a positive correlation holding infant mortality constant.

The most empirically testable work in this area is that of Rose Frisch. She has shown that the onset of menarche is determined largely by the attainment of a minimal fat/weight ratio (Frisch et. al. 1973) and that after menarche, there is a threshold level of the fat/weight ratio below which ovulation ceases (Frisch and McArthur 1974). The model inherent in her work on post-menarche fertility is relevant to this research. This work suggests nutrition has a strong relation to the frequency of anovulatory cycles. The level of the fat-weight ratio which leads to a cessation of ovulation is posited to be well above the starvation level. If so, it is possible that a substantial proportion of a female population with a relatively low calorie level, but otherwise generally healthy, could be below the fat-weight ratio threshold. Frisch's work also suggests that at low effective calorie levels, even above the threshold level, all of the relevant intervals (waiting time to conception, interval to a successful live birth, post-partum non-susceptible period) are lengthened. Frisch's results are based primarily on clinical experiments involving only a small, non-representative sample of human subjects; thus, their implications for the nutrition-fertility relationship in an actual population are highly controversial.[1]

The main purpose of the present paper is to examine the crude relationship between different measures of household nutrition and fertility in a less-developed country, Zaïre, in which there is substantial malnutrition [Johnson (1975)]. The main emphasis will be on the total effect of these nutrition measures on the probability of later births after the first live birth when infant mortality rates are held fixed. In this paper, we investigate the relationship between nutrition and fertility through examination of closed birth intervals, including the average interbirth interval and the length of the first and second closed birth intervals. We also examine the probability that the second birth occurs at any given length of time after the first birth, whether or not the second birth occurred before the survey date. Both types of analysis give some support to the existence of a negative relationship between calorie consumption and the length of time between births. Our strongest result is that there is a significant direct association between the monthly probability of conception and one nutrition measure, purchased total calories per adult equivalent, among households with no infant deaths in the 20-24 year old female cohort. This result is subject to at least two interpretations: (1) a joint set of factors causes some households to choose at once greater abstinence during lactation, longer lactation periods, and fewer purchased calories per adult equivalent; and (2) there is the positive correlation between a woman's fat-weight ratio and her fecundity which Frisch observes. Further empirical tests seem to support the second interpretation. The available data for Zaïre indicate a weak positive rather than negative association between purchased calories and lactation period length. Furthermore, purchased calories per adult equivalent remain statistically significant even after controlling for some exogenous variables which may influence both purchased nutrition and fertility decisions.

I. Closed Birth Intervals for Married Women

The most direct method to investigate the relationship of nutrition to marital fertility is to study fertility unaffected by infant deaths. For this purpose, we analyzed households in the 1969 socio-economic survey of Kinshasa headed by a male with only one wife, where the wife was aged 20-24 at the survey date. Since we knew neither the birth nor death dates of dead children but only how many children of the wife had died, we limited this analysis to women for whom all children were alive and present in the household at the time of the survey Restricting the analysis to households in which there have been no infant deaths is a common practice in studies of birth intervals, in order to remove the effect of mortality on the length of the non-susceptible period [D'Souza (1974)].

Each household in the survey was visited once a week for four weeks. The amount of money spent on each commodity by commodity was recorded. Calorie and price tables [Houyoux (1973)]were used to calculate the total number of calories and grams of protein purchased in the month. Since for each household, the sex and month and year of birth (when available) were recorded, the number of adult equivalents in each household could be calculated, and then the number of purchased calories and grams of protein per adult equivalent per day by household could be found. These are the nutritional variables used in this study.

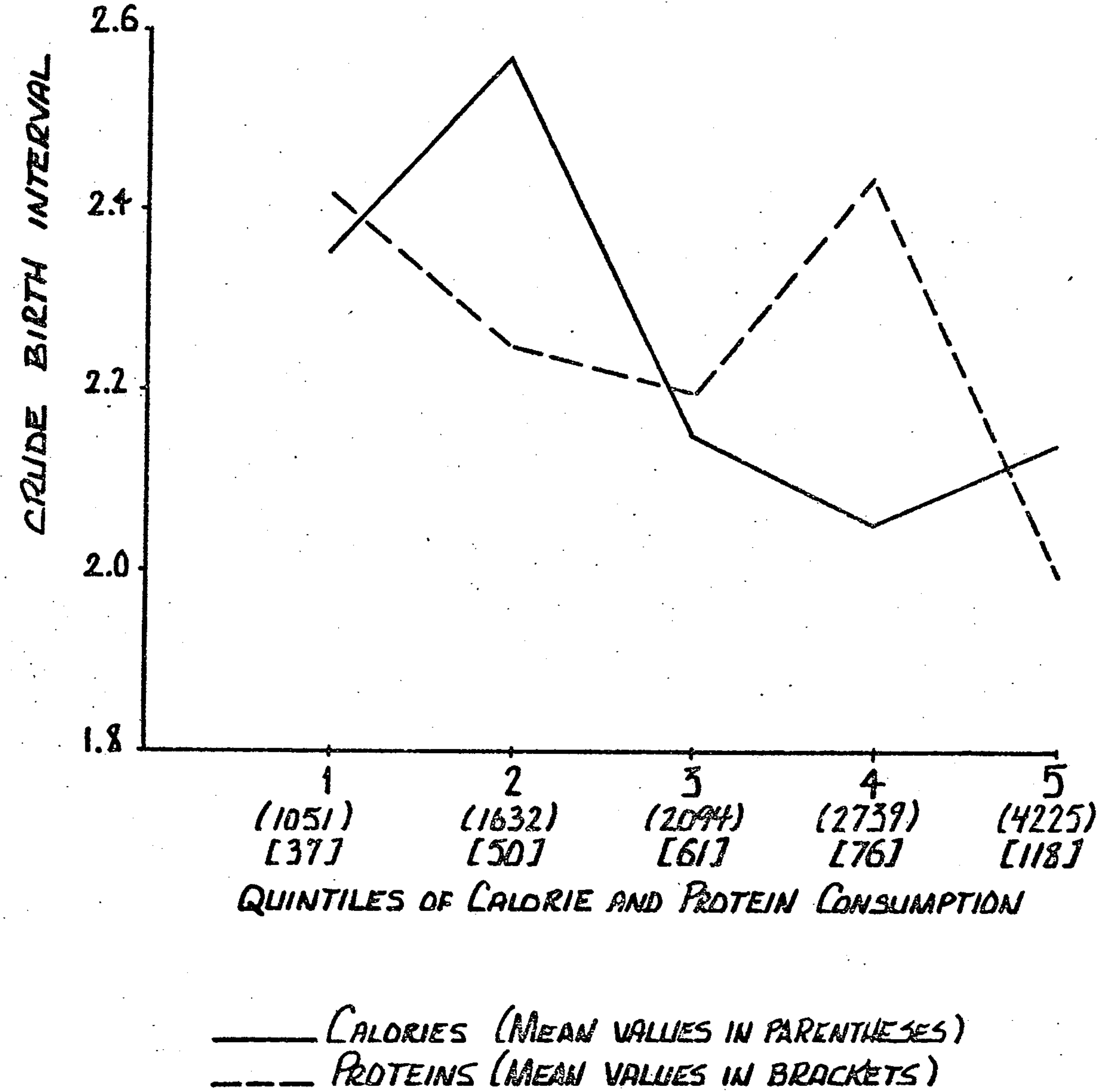
The implicit assumptions are: 1) purchased calories and grams of protein in a time period are strongly positively related both to calories and proteins consumed in that time period and to calorie and protein consumption at earlier times and 2) calories and grams of protein are dispensed within households proportionately to the number of adult equivalents which each person represents.

Our analysis is based entirely on birth interval data for the cohort of married women aged 20-24. It is well known that (a) the length of the interbirth interval and (b) the degree of memory-lapse bias each increase as the age of the mother increases. If the latter effect is present, some households will be included in the sample simply because they forgot all of their dead children; in these households, intervals between children ever born (estimated from the ages of surviving children) may well be overstated. Another reason for restricting the analysis to a relatively narrow age group of younger women is that current nutrition for these women is more closely related to nutrition during childbearing than it would be for older women. Finally, the proportion of households eliminated because of infant death is smaller in the cohort aged 20-24 (10%) than in older cohorts (22% for the entire sample).

The crude birth interval was estimated by subtracting the date of birth of the youngest from the date of birth of the oldest child and dividing by the total number of children minus 1. Figure 1 shows this average length of the interbirth interval for households where the wife was aged 20-24 at the time of the survey by values of the quintiles of the consumption of calories and grams of protein per adult equivalent per day for this age group. The values are shown by quintiles in order to minimize the effect of extreme and probably erroneous values of protein or calorie consumption. If the hypothesis indicating a pro-natalist effect of food consumption is correct, there should be a negative relationship between food consumption and the average length of the interbirth interval. This generally is shown for calorie consumption, while the results for protein consumption are very mixed. Mean values of calories are shown in parentheses and of proteins in brackets. The quintiles are for households appearing in the calculations. This is consistent with Frisch's hypothesis since calorie intake should be more closely related to the fat-weight ratio than protein intake. However, there is also the behavioral interpretation that hunger may be more sensitive to variation in calorie than it is to variation in protein intake. To the extent that hunger is a major determinant of the relative length of lactation periods, birth intervals will tend to be more responsive to relative differences in protein intake. The bulk of the rest of the analysis centers on calories per adult equivalent.

FIGURE 1

RELATIONSHIPS OF CALORIE AND PROTEIN CONSUMPTION
TO THE CRUDE BIRTH INTERVAL FOR WOMEN
AGED 20-24, KINSHASA 1969



<u>Regression Model</u> The statistical significance of the inverse association between actual calories per adult equivalent and the length of closed birth interva is determined by means of regression analysis. Denote calories per adult equivalent for the household at the time of the survey by C, education of the husband by EH, education of the wife by EW and the age of the oldest child by AGE. Let I represent the length of the relevant interval between live births measured in years. Then the regression equation to be estimated may be written in the form

$$(1) \qquad \sqrt{I-.75} = a_0 + a_1\ C + a_2\ EW + a_3\ EH + a_4\ AGE + \eta$$

where η is an error term, assumed normally distributed. The constant .75, is subtracted from the interval between live births to give an estimate of the interval between the earlier birth and the next conception which resulted in a live birth, which is followed by a gestation period of .75 years to the later birth. The square root transform of the waiting time to conception appears as the dependent variable rather than the variable itself because the distribution of waiting times to conception is skewed to the right in most populations. Standard practice among demographers is to assume that this distribution is approximated by a convolution of a normal and an exponential density function. D'Souza (1974) has shown, using othe data, that the square root of the estimated waiting time to conception is almost normally distributed. The chi-square goodness of fit test indicates that the distribution of the waiting time to conception in the Kinshasa data set is significantly different from the normal at the 5% level, but that the square root transformation at the same confidence level is not.

The coefficients in the regression equations have the following hypothesized signs

$a_1 < 0,\ a_2 < 0,\ a_3 < 0,\ a_4 > 0.$

The hypothesized signs of the coefficients for the education variables are generally ambiguous. If women in the cohort aged 20-24 are practicing birth control, other than observing the usual intercourse taboos during lactation, then a solid economic argument may be made for the coefficients, a_2 and a_3, being positive if substitution outweigh income effects [Ben-Porath (1973) and Willis (1973)]. This argument centers around hypothesized positive associations among education, potential wages, and the price of children in the household production model. But such an effect resulting from an increase in the price of children may be insignificant if, as assumed, the couple is still highly uncertain whether it can attain its desired number of surviving children because of subfecundity and high infant mortality. In this case, the couple may not wish to control its post-lactation fertility at all, particularly when the wife is i her early reproductive years. Although demand effects cannot be totally ruled out a priori, the sign of the coefficients a_2 and a_3 are presumed to depend on the qualitative relation between wife's education and fecundity. All available evidence indicates that the women in the cohort age 20 to 24 are not practicing birth control other than intercourse taboos during lactation.[2] The qualitative effect of education on fecundity is hypothsized to be positive, which implies negative values for a_2 and a_3. The education of the husband may be a better indicator of average nutrition of the household over its reproductive life than its calorie consumption at the time of the survey. Wife's education may be a proxy for her nutritional level and overall health prior to marriage.

The age of the oldest child at the time of the sample, AGE, represents an upper bound on the waiting time to conception. The expected value of the interval length is, therefore, conditioned on this variable. Take, for example, the expected value of the interval between the birth of the first child and the conception of the second for a household whose oldest child was age a + .75 at the time of the sample. This is given by the expression

$$E(S_1|\text{age} - .75 = a) = \frac{\int_{S=0}^{a} S_1 f(S_1|X)ds_1}{\int_{S=0}^{a} f(S_1|X)ds_1}$$

where $f(S_1|X)$ is the probability that the second child will be conceived exactly S years after the birth of the first and X is a vector of independent variables determining the probability of conception. Assuming that the probability density function is the same for higher intervals, the expression for the expected value of the next interval is

$$E(S_2|\text{age} - 1.50 = a) = \frac{\int_{S_2=0}^{a} S_2\, f(S_2|X)F(a-1.50-S_2|X)ds_2}{\int_{S_2=0}^{a} f(S_2|X)\, F(a-1.50-S_2|X)ds_2}$$

where F(v) is the distribution function associated with the density function f(S), i.e., $F(v) = \int_{S=0}^{v} f(S)\, ds$. It is easily shown that both $E(S_1 \mid \text{age} - .75 = a)$ and $E(S_2|\text{age}-1.50 = a)$ are increasing functions of age. Hence, the hypothsized sign of a_4 is positive.

Before turning to the regressions themselves we should note that the variable, calories per adult equivalent, may be endogenous rather than exogenous as assumed. If this is true, the coefficient for this variable will be biased. However, under most reasonable assumptions, the bias will be against our hypotheses. There are two main ways in which calories per adult equivalent may be interpreted as an endogenous variable. In the first case, suppose that purchased calories are constant across households and the number of children ever born randomly varies for households in which the age of the oldest child is the same. Then under these assumptions it is clear that there w be a positive association between the length of any given birth interval and calories per adult equivalent. The more children a household has, the shorter the birth interval and the smaller will be the number of calories per adult equivalent. The number of adult equivalents involves a weighted sum of the number of children, as well as other family members. This effect will bias the coefficient a_1 in a positive direction, against our hypothesis.

Secondly, omitted variables may influence purchased calories as well as the birth interval. Households in which the wife has a relatively long lactation period would cet. par. be expected to have a higher level of purchased calories and longer birth intervals unless lactating mothers consume fewer additional calories than they produce in the form of milk. To the extent that there is a positive association between purchased calories per adult equivalent and the length of the lactation period, the coefficient a_1 will be biased against our hypothesis. However, it should be recognized that the cost per calories is considerably higher for a human milk substitute than it is for breast milk [Berg (1973)]. The main determinant of expenditure on human milk substitutes, then, may be income. If human milk substitutes are considered normal goods by the household and make child rearing less intensive in the wife's time, we may expect their quantity demanded to increase with the head of household's income and purchased calories. The latter would generally lead to increases in the value of wife's time and substitution of commodity for time inputs in child rearing. By this line of reasoning, households with higher levels of purchased calories per adult equivalent will have shorter lactation periods and the coefficient a_1 will be biased in favor of our hypothesis. However, there appears to be a weak positive correlation between lactation period length and purchased calories per adult equivalent.[3]

There is also a measurement problem with the calorie variable. It is known that at low levels of calorie consumption, proteins are utilized as calories rather than to maintain nitrogen balance [FAO-WHO (1973)]. If the contribution of proteins to effective calorie consumption had been considered, the result would have been to raise the measured calorie consumption of the low calorie part of the sample. We know of no way to properly convert protein consumption into effective calorie consumption at varying levels of calorie consumption.

Parameter Estimates We experimented with several different birth intervals in the regression analysis, including the crude birth interval as previously defined and the first and second closed birth intervals. Separate regressions were run in which the square root of each of the intervals minus .75 years appeare as the dependent variable. Another set of regressions was run on pooled data in which the unit of analysis is each interval minus .75 years rather than the household. In these regressions, separate dummy variables were included which correspond to the order of the birth which initiated the interval. When controlling for education, the calorie coefficient was negative but insignificant at the 5% percent level except in the regressions involving the pooled interval data set. The same regressions were run on a more homogeneous data set in which the households were constrained to have at least three children (See Table 1). These regressions weight the households with relatively long intervals more evenly within the sample than did those based on the larger data set. In this case, the coefficients of calories per adult equivalent has the hypothesized negative sign and are significantly different than zero at the 5% level except in the case where wife's education is included. The latter specification is not acceptable by conventional statistical standards since the t-ratio on wife's education is below 1.67. Even when wife's education was included, the t statistic of calories was significant at the 5% level, except for the second interval where the t was 1.99.

Table 1

Regressions of Birth Intervals of 20-24 Year Old Women

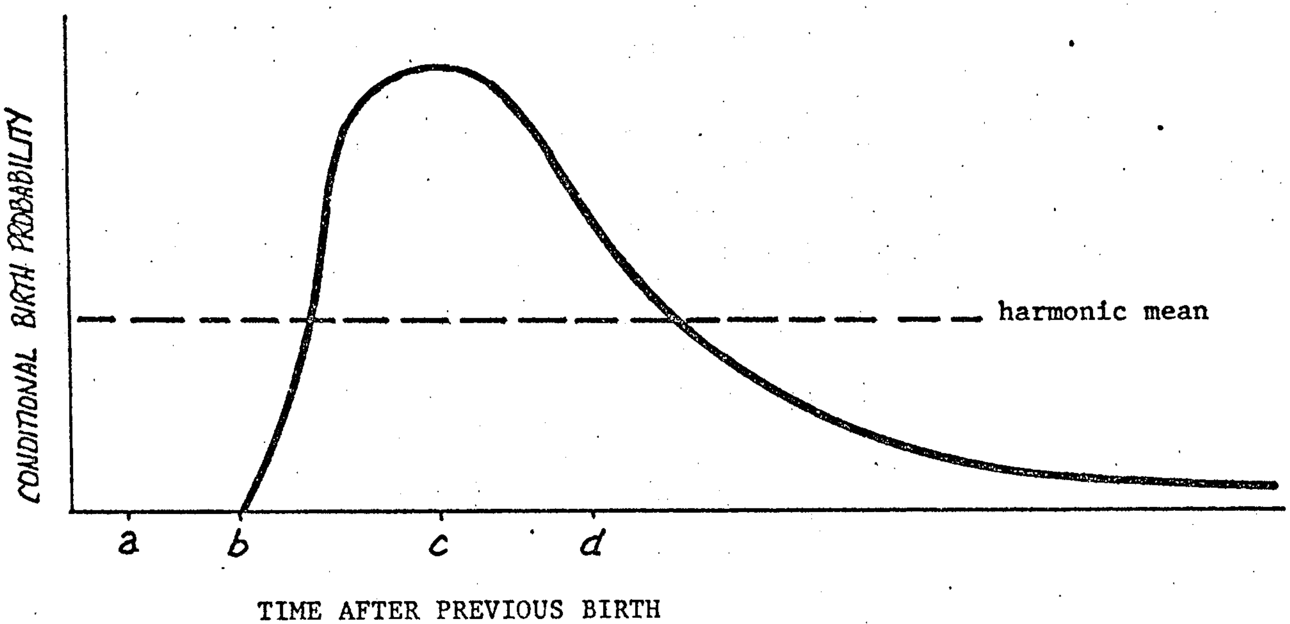
Interval	First	First	Average	Average	Second	Second
Calories	-5.4×10^{-5} (1.63)	-4.6×10^{-5} (1.37)	-6.1×10^{-5} (3.06)***	-4.6×10^{-5} (2.66)***	-7.0×10^{-5} (2.56)**	-5.6×10^{-5} (2.15)**
Age of Oldest Child	5.3×10^{-2} (2.10)**	5.4×10^{-2} (2.18)**	7.1×10^{-2} (4.71)***	7.4×10^{-2} (5.75)***	5.0×10^{-2} (2.36)**	5.3×10^{-2} (2.69)***
Husband's Years of Education		-1.5×10^{-2} (1.32)		-2.8×10^{-2} (4.82)***		-2.7×10^{-2} (3.07)***
Constant	.94	1.02	.85	1.00	1.02	1.17
R^2	.106	.132	.346	.535	.175	.296
F	3.43	2.90	15.31	21.88	5.95	7.71
n	61	61	61	61	59	59

There are several problems with this analysis. To begin with, restricting a sample even to households with at least two children precludes us from testing the effectiveness of a model explaining the fertility behavior of relatively sub-fecund groups. Secondly, as Heckman and Willis (1975) have shown the reciprocal of the expected value of a birth interval may imply an estimate of the expected monthly probability of conception substantially below the true expected monthly probability of conception, during the post-partum period. This is immediately clear from figure 2. For any group of households having the same level of nutrition, the conditional probability of conception (given no conceptions since the birth of the first child) will rise initially as more women become susceptible. Then it will fall as the relatively fecund households leave the sample through conception. The reciprocal of the average of the birth intervals (or their square roots) represents the harmonic mean of the conditional conception probabilities (or their square roots). For the harmonic mean to be an unbiased estimator of the expected value of a random variable, the variable must have a geometric distribution such as that illustrated by the dotted line in figure 2. The greater the variation in monthly conception probabilities due to differences in fecundity not explained by nutrition, the more their harmonic means will be biased toward zero. This error in the dependent variable may cause the t-statistics for calories per adult equivalent and other independent variables to be biased downward , particularly in a sample where there is large variation in fertility across households [Heckman and Willis (1975)].

Table 1 (continued)

Regressions of Birth Intervals of 20-24 Year Old Women with at Least Three Children

Interval	All	All	All
Calories	-6.8×10^{-5} (3.92)***	-5.5×10^{-5} (3.19)***	-5.5×10^{-5} (3.08)***
Age of Oldest Child	4.8×10^{-2} (3.45)***	5.2×10^{-2} (3.79)***	5.3×10^{-2} (3.60)***
Husband's Years of Education		-1.9×10^{-2} (3.10)***	-1.9×10^{-2} (2.99)***
Birth Order			-1.9×10^{-2} (0.14)
Dummies First			1.2×10^{-2} (0.09)
Second Third			-2.4×10^{-2} (0.16)
Constant	1.01	1.10	1.10
R^2	.155	.211	.214
F	12.45	12.03	5.99
n	139	139	139



Typical Conditional Birth Probabilities

Figure 2

This seems to be a characteristic of the first closed interval data set where households with only two births are included and may explain why the coefficient for the calorie per adult equivalent variable was statistically insignificant in this case.

II Open Interval Estimates

In this section we examine a data set which has been expanded to include couples having only one child. We concentrate on directly explaining the conditional probability of conception after the first birth using this expanded data set. For each household in our sample, point a in Figure 2 is the birth of the first child. Since births rather than conceptions are observed, the earliest time at which a second birth could occur is nine months after a, at b. Six month periods of household experience after b are our units of analysis. If a household experiences its second birth in period i after b, then it is excluded from the set of households under observation in period (i + 1). Also, a household will not be observed in period (i + 1) if the survey date arrived first, i.e. if the first child was born less than (9 + 6(i + 1)) months before the survey.

Table 2 shows the number of households under observation in each six month period after b and the probability that the second birth occurred in each period (second births in period/households in period). These probabilities are graphed in Figure 3. As expected, these conditional probabilities exhibit the general pattern shown in Figure 2, with fluctuations in late periods when the number of households under observation becomes small. By the end of the eleventh period (75 months after the first birth), 72% of the original 116 households had experienced the second birth while under observation. Thus it was far more likely that a household left the sample through the occurrence of the second birth than through truncation due to the arrival of the survey date.

Frisch and McArthur (1974) have contended that the primary effect of nutrition on fertility is through differences in the waiting time to conception after the cessation of lactation amenorrhea rather than through differences in the length of the amenorrhea.

Table 2

Households Under Observation and Second Birth Probabilities in Six Month Time Periods Beginning Nine Months After the First Birth

Period	Months After First Birth	Number of Households Under Observation	Second Birth Probability
1	9-14	116	.026
2	15-20	96	.167
3	21-26	69	.420
4	27-32	39	.538
5	33-38	18	.222
6	39-44	14	.214
7	45-50	11	.273
8	51-57	8	.125
9	58-63	5	.200
10	64-69	4	.250
11	70-75	3	.333

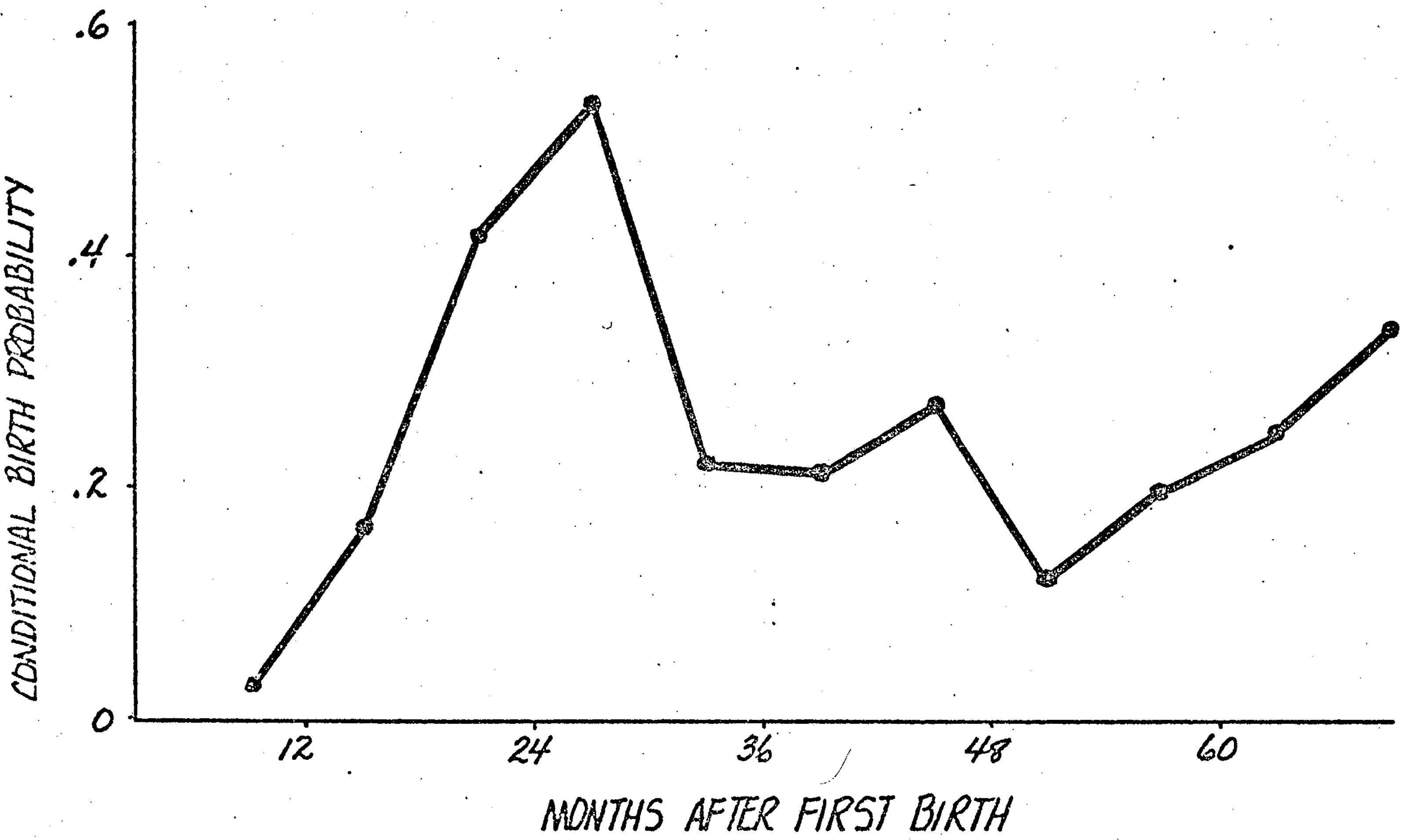


Figure 3

Observed Conditional Birth Probabilities

An appropriate functional form for the threshold relationship posited between calories and conception probability after lactation amenorrhea is the logistic [Nerlove and Press (1973)]. Thus a logistic equation of the following form was estimated :

$$(2.1) \qquad Pr(B) = 1/(1 + e^{-V})$$

where $V = a_1 Cal_1 + a_2 Cal_2 + a_3 Dum_1 + \ldots + a_{13} Dum_{11}$

In the above equation, $Pr(B)$ is the probability that the second birth occurs to a household in the period under observation. Dum_i assumes the value one if the period of observation is period i and the value zero otherwise. Cal_1 is calories per adult equivalent if Dum_1 is one and is zero otherwise. Cal_2 is zero if it is the first six-month time period and is calories per adult equivalent if the period of observation is not the first period.

A model involving only calories per adult equivalent as an independent variable may be subject to serial correlation due to omitted variables (for example, permanent post-partum sterility) specific to individual households but invariant over time. The use of a dummy variable for each period is a method of correcting for such serial correlation which changes the composition of the sample across time periods. If the effect of the time period is to shift the location but not the scale of the logistic density functio the dummy variable procedure will yield unbiased estimates. The effect of calories in the first period is considered separately from all other periods, since most households are still in post-partum amenorrhea at that time, and calories are not expected to be strongly related to whether a household experiences an exceptionally short amenorrhea period.

The coefficients in (2.1) were estimated using a maximum likelihood procedure [Nerlove and Press (1973)]. Probabilities of birth in a given period are not observed but rather whether or not a birth occurs. A maximum likelihood procedure is appropriate for a logistic specification such as ours, where the dependent variable only assumes values of zero or one and some of the independent variables are continuous. Such a model conforms neither to the assumptions of linear models, such as linear regression, nor to the requirements of contingency table analysis.

Table 3 shows the estimated coefficients, along with their asymptotic t-values and the elasticities of the calorie variables. Specifications are shown with only the dummy variables used as independent variables and with Cal_1 and Cal_2 added stepwise. As expected, calories in the first period, while most households are in lactation amenorrhea, are not significantly positively related to the occurrence of a second birth. The results of the log likelihood ratio test show that the addition of Cal_2 to the estimation equation significantly improves the fit of the model to the data at the five per cent level, while the addition of Cal_1 to the model which only includes the dummy variables, does not improve the fit.

The elasticities show the percentage change in the dependent variable which would result from a one percent change in the independent variable, at the mean of the independent and dependent variables. The effect is quite strong for Cal_2. Another way of interpreting the effect of calories on second birth probabilities is shown in Figure 4. The cumulative second birth probabilities predicted by the third specification in Table 3 are shown at calorie levels one standard deviation above (3741) and one standard deviation below (1047) the mean calories per adult equivalent level of the 116 households in period 1. At the end of period 4 (33 months after the birth of the first child), 88% of households at the higher but only 70% of households at the lower calorie level would have experienced the second birth.

The calories per adult equivalent variable has much greater statistical significance in this birth probability analysis than it did when the first closed interval was the dependent variable. Moreover, it is important to note that when the dummy variables providing information on conceptions in specific previous periods were removed, the estimated coefficients and t-ratios for the calories per adult equivalent variables fall substantially. Hence, failure to correct for serial correlation in this context may well have lead to coefficients biased toward zero in the case of the calories per adult equivalent variable.[4]

Table 3

Maximum Likelihood Logit Estimates of Second Birth Probabilities

Variable	(1)	(2)	(3)
Cal_1 Calories in period 1		9.68×10^{-5} (0.24) [.065]	9.68×10^{-5} (0.24) [.067]
Cal_2 Calories in periods 2-11			2.89×10^{-4} (2.35)* [.410]
Dum_1	-3.63 (6.20)**	-3.87 (3.29)**	-3.87 (3.29)**
Dum_2	-1.61 (5.88)**	-1.61 (5.88)**	-2.32 (5.46)**
Dum_3	-.322 (1.32)	-.322 (1.32)	-.962 (2.61)**
Dum_4	.154 (0.48)	.154 (0.48)	-.390 (0.98)
Dum_5	-1.25 (2.21)*	-1.25 (2.21)*	-1.78 (2.90)**
Dum_6	-1.30 (1.99)*	-1.30 (1.99)*	-1.88 (2.67)**
Dum_7	-.981 (1.45)	-.981 (1.45)	-1.54 (2.13)*
Dum_8	-1.95 (1.82)	-1.95 (1.82)	-2.54 (2.30)*
Dum_9	-1.39 (1.24)	-1.39 (1.24)	-2.00 (1.72)
Dum_{10}	-1.10 (0.95)	-1.10 (0.95)	-1.79 (1.49)
Dum_{11}	-.693 (0.57)	-.693 (0.57)	-1.31 (1.03)
	72.5**	72.5**	78.1**
Log-likelihood function	-164.0	-163.9	-161.2
Log-likelihood ratio test	.0562		5.550*

Coefficients are given first. Two tailed t-values are in parentheses.
Elasticities are in brackets.
*p < .05, **p < .01

Predicted Cumulative Birth Probabilities

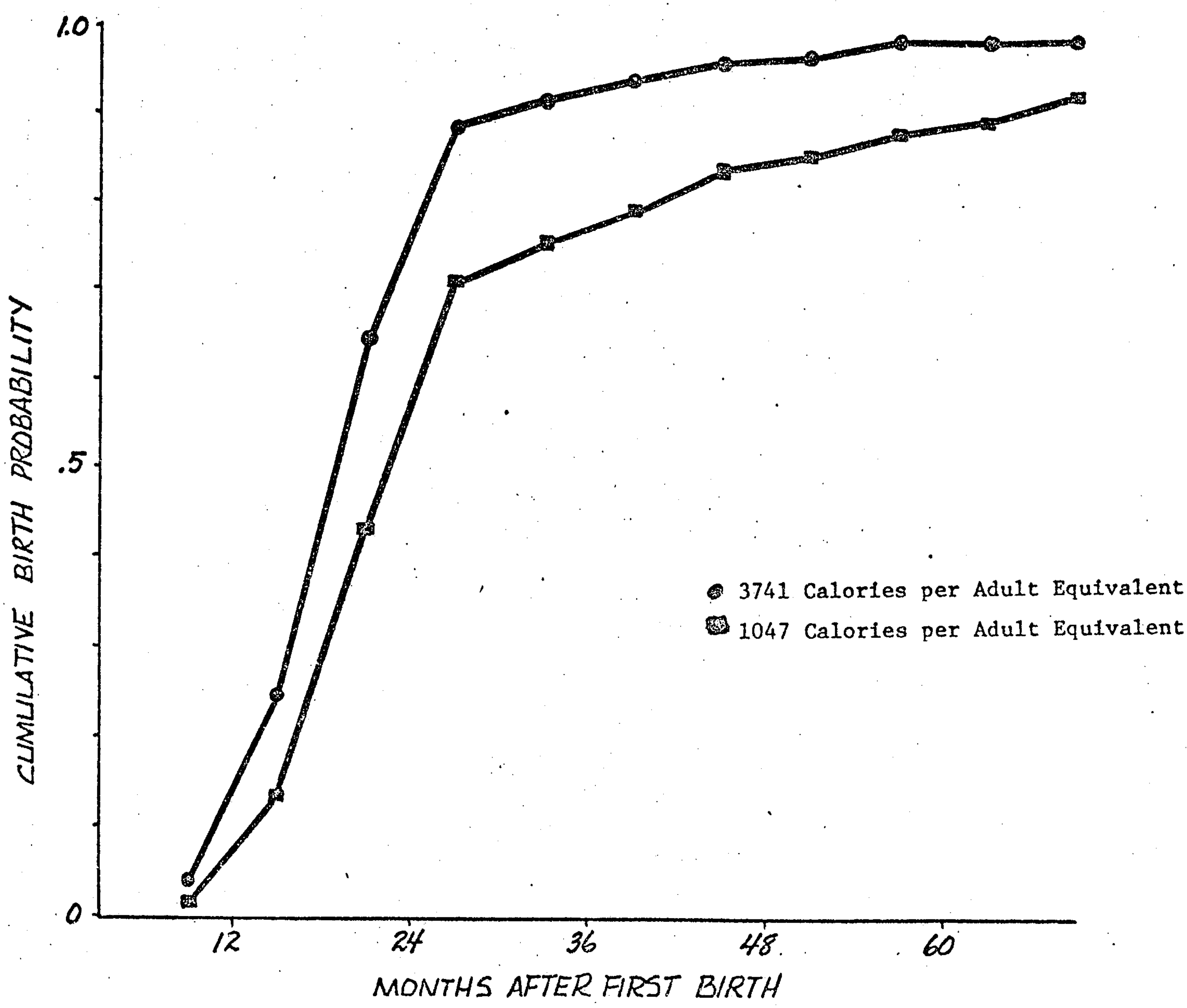


Figure 4

Experiments were performed with a number of economic variables which may jointly influence nutrition and fertility choices, including husband's and wife's education, total household income per adult equivalent and total household consumption per adult equivalent. The statistical significance of the calories per adult equivalent variable survived the inclusion of these additional variables in all cases. Except for total income which was endogenous, due to the inclusion of wife's income, the socio-economic variables were not statistically significant.

What may be happening is that socio-economic variables may influence decisions about the calorie content and other aspects of the nutritional adequacy of the diet, while the nutritional adequacy affects fertility. Thus, while socio-economic variables may be important for fertility in Zaire, their effect on fertility may be primarily indirect through nutrition.

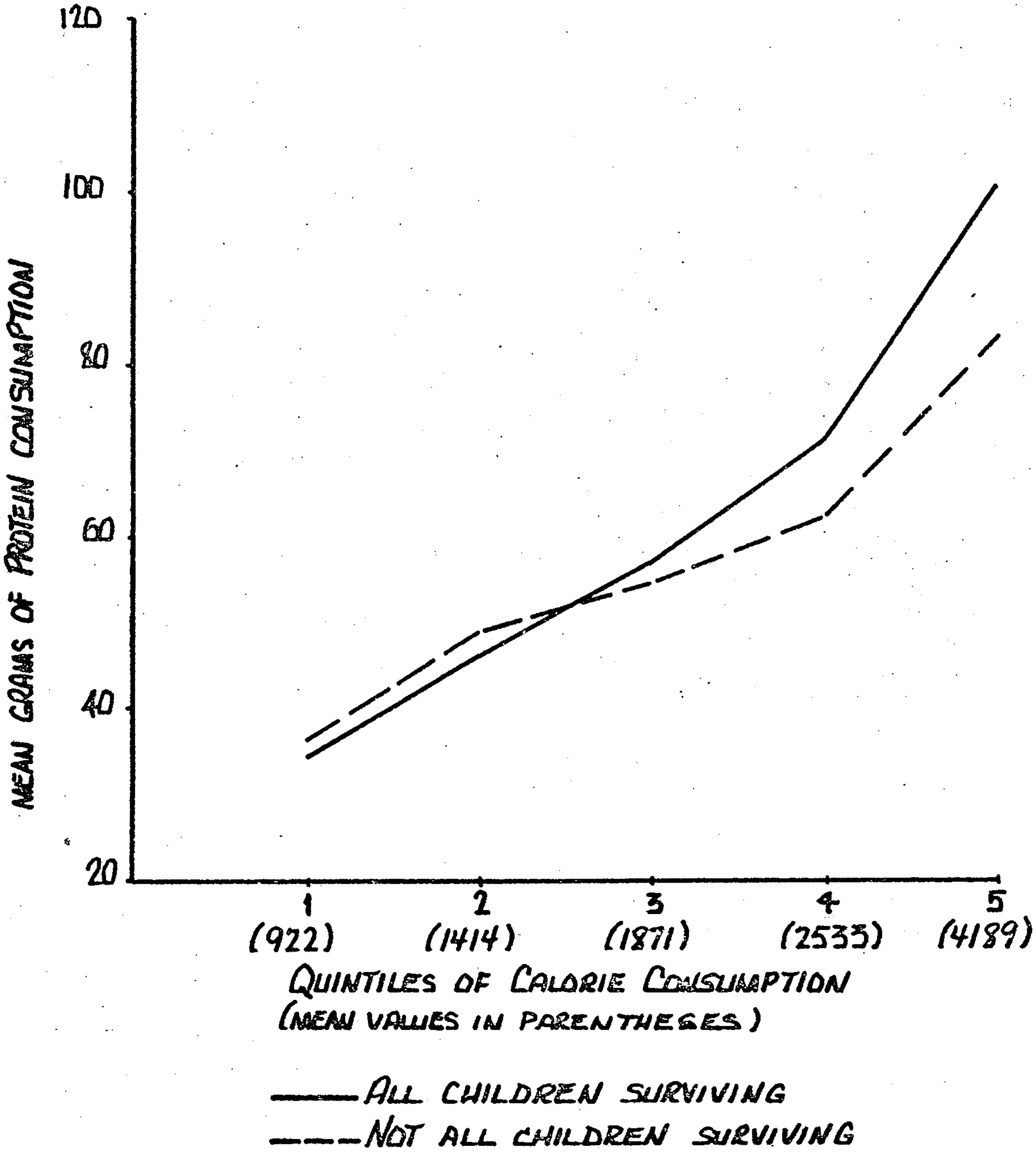
III. Policy Implications

The main policy implication of this investigation is that rises in calorie consumption p adult equivalent associated with the early stages of modernization may be expected to increase fertility in non-contracepting populations with no change in infant mortality rates. If, however, infant mortality also declines, the total effect of the increase in calories per adult equivalent on the fertility of women in household experiencing infant deaths will be ambiguous.

There is some evidence that if calorie consumption per adult equivalent can be held constant and protein consumption increased both infant mortality and fertility may fall. As already indicated, at low levels of calorie consumption, proteins are not used primarily as proteins to maintain nitrogen balance but rather simply as additional calories [(FAO-WHO,1973)]. Figure 5 illustrates the point that, in the Kinshasa data set, infant mortality may depend on the amount of protein, not calories consumed, once a critical minimum level of calorie consumption has been attained. The values in Figure 5 are averages across each five year age group of women aged 25-39, who had borne at least one child. The quintiles are for households with one wife age 15 or older. Below age 25, there were very few dead children. Above age 40, the retrospective fertility data

FIGURE 5

MEAN PROTEIN CONSUMPTION BY CALORIE CONSUMPTION AND BY SURVIVAL OF CHILDREN



become somewhat suspect. A t-test for the highest quintile of calories shows that women with no dead children had significantly higher protein consumption than women with same dead children at the 5% level, one-tailed test. Furthermore, there is evidence for the 20-24 female cohort that calories per adult equivalent had a statistically significant association with different fertility measures and that proteins did not. This was true both for separate regressions in which only one of these nutrition measures appeare as an independent variable and also for a single regression in which both appeared. Therefore the available evidence is consistent with arguments that policies designed to eliminate protein deficiency will have an anti-natalist effect provided calorie consumption is held fixed above a critical minimum level.[5] This effect would be indirect through induced declines in infant mortality. Note that statistical results for other countries indicate that infant survival rates have a strong negative association with the fertility of younger as well as older mothers in poorly nourished populations with long lactation periods [D'Souza (1974) and Preston (1975)].

The results presented certainly are consistent with a positive biological relationship between calorie consumption and fertility. However, a more general way of analyzing fertility determination is through a system in which nutrition, lactation-related behavior and fertility are simultaneously determined. Socio-economic variables are likely to influence nutritional and lactation decisions and thus influence fertility. In order to properly identify such a system, variables influencing food consumption but not fertility are needed. For instance, one could use commodity price differences to s the response of households' nutritional decisions to differing market conditions. One could only do this for a sample in which different households are facing different prices, a phenomenon which is absent in cross-section data for a single city. The combined examination of similar data from a number of Zairois cities may allow more to be discovered about the complex biological and behavioral determinants of fertility than has yet been possible.

Footnotes

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[1]Frisch's sample is made up primarily of patients with anorexia nervosa, which may in itself lead to cessation of ovulation. Earlier studies show a direct relation between nutrition anovulation during the post-menarche period but involve subjects under severe psychological stress. See Sydenham (1946) and Smith (1947).

[2]Coale and Trussell (1974) have developed a summary measure of voluntary fertility control, m. This measure represents the degree of deviation in the observed age-specific marital fertility schedule from the natural fertility schedule. The higher the negative slope of the observed age specific fertility schedule within a five year age cohort over age 25 relative to that of the natural fertility schedule, the higher will be the value of m. We estimated m for five year cohorts of women aged 25 to 39, by separate education groups of husband and wife from the 1967 socio-demographic survey data. The coverage of this survey is 10 times greater than that of the 1969 socio-economic survey of the same city. The estimates of m were all quite close to zero and were negative except for the cohort aged 35 to 39 at the highest educational levels. Even so, the highest positive value was .15 for the group with secondary educated husbands, which still indicates a virtual absence of voluntary fertility control.

[3]Data on lactation periods were obtained from the 1974 Measles Vaccination Survey of Kinshasa. We grouped data by communes (22 in number) for comparison with the 1969 Socio-Economic Survey. Regressions of lactation period means against calories per adult equivalent means for each commune showed an insignifican positive correlation, excluding the richest commune (8 cases out of 2089) in the Measles Survey. Coefficients for commune dummy variables were obtained by regressing individual child lactation periods against age of mother at birth and commune dummy variables. These dummy variables were jointly significant at the 1% level, and several were individually significant at the 10 percent level. Regressions with the coefficient estimates of these communes as dependent variables yielded coefficients that were positively, but insignificantly, correlated with mean calories per adult equivalent by commune.

For a description of the Measles Vaccination survey, see Adelman, (1975).

[4]Although fat-weight ratios may have a threshold relationship to fertility, there is no conclusive evidence that calories per adult equivalent have such an effect on birth probabilities. A linear probability formulation

$$Pr(B) = b_2 \, Cal_1 + b_2 \, Cal_2 + b_3 \, Dum + \ldots + b_{13} \, Dum_{11}$$

was also estimated using ordinary least squares. The estimate of the coefficient b_2 was positive and significant at the five percent level, which was no worse than the result obtained from the logistic specification.

[5]Extreme protein malnutrition, as well as calorie malnutrition, may be one cause of the low total marital fertility in Kisingani as compared with Kinshasa. See Houyoux (1972).

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