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Predictors of Fecundability and Conception Waits Among the Dogon of Mali

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ABSTRACT Surprisingly little is known about the mechanisms that underlie variation in female fertility in humans. Data on this topic are nonetheless vital to a number of pragmatic and theoretical enterprises, including population planning, infertility treatment and prevention, and evolutionary ecology. Here we study female fertility by focusing on one component of the interbirth interval: the waiting time to conception during menstrual cycling. Our study population is a Dogon village of 460 people in Mali, West Africa. This population is pronatalist and noncontracepting. In accordance with animist beliefs, the women spend five nights sleeping at a menstrual hut during menses. By censusing the women present at the menstrual huts in the study village on each of 736 consecutive nights, we were able to monitor women's conception waits prospectively. Hormonal profiles confirm the accuracy of the data on conception waits obtained from the menstrual hut census (Strassmann [1996], Behavioral Ecology *7:*304–315). Using survival analysis, we identified significant predictors of the waiting time to conception: wife's age (years), husband's age (<35, 35–49, >49 years), marital duration (years), gravidity (number of prior pregnancies), and breast-feeding status. Additional variables were not significant, including duration of postpartum amenorrhea, sex of the last child, nutritional status, economic status, polygyny, and marital status (fiancée vs. married). We fit both continuous and discrete time survival models, but the former appeared to be a better choice for these data. Am J Phys Anthropol 105:167–184, 1998.

An improved understanding of human evolutionary ecology will require clarification of the causes of differential reproduction during our evolutionary past. In particular, it will require clarification of the evolved mechanisms that underlie differences in female fertility. Because fossil remains are of limited value for this purpose, the best alternative is to study natural fertility populations. In these populations, the effect of distal variables such as land ownership and proximal variables such as nutritional status can be tested for an effect on fecundity without the confounding effect of contraception.

Most of the relevant research has been conducted by demographers, who have made two major contributions. First, they have identified many variables that have a direct effect on fertility and through which the effects of distal variables are mediated. This is known as the proximate determinants approach (Bongaarts and Potter, 1983). Sec-

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ond, they have recognized that to understand fertility it is useful to disaggregate the birth interval into separate components: postpartum amenorrhea, the waiting time to conception among cycling women, pregnancy, and the time added by fetal loss. Here we use both discrete and continuous time survival analysis (Lawless, 1982) to identify the independent variables associated with the waiting time to conception in a Dogon village in Mali, West Africa. We also estimate the impact of the covariates on effective fecundability, defined as the monthly probability (continuous time model) or cyclewise probability (discrete model) of a conception that results in a live birth (Wood, 1994).

The Dogon are a natural fertility population (sensu Henry, 1961) with a total fertility rate of 8.6 live births per woman. This high fertility rate is driven in part by a high mortality rate (20% mortality by age 1 year, 46% by age 5 years, N = 388) (Strassmann, 1992). Dogon villages are particularly suitable for studying variance in conception waits because the confounding effect of contraception is completely absent. Most previous studies of conception waits were conducted in populations that were not uniform with respect to contraception use. Some of these studies omitted the contracepting women from the sample and estimated fecundability from the remaining women. If the contracepting women had a higher probability of conception, this procedure underestimated fecundability (Goldman et al., 1985; Larsen and Vaupel, 1993; Baird et al., 1994). In some populations, fecundability is depressed not by contraception but by sterility caused by venereal disease. In our study village, however, the prevalence of primary sterility (no pregnancy ever) was only 0.9% (1 out of 113 women). The prevalence of secondary sterility (after the birth of one child) was only 1.8% (2 out of 113 women). Nor was there evidence of AIDS at the time of the fieldwork reported here. Thus, we were able to measure fecundability in the absence of these major confounding variables.

Another advantage of the Dogon as a study population derives from their menstrual taboos. During menses Dogon women spend five nights sleeping at a menstrual hut, which made it possible to monitor female reproductive status prospectively without interviews. It was readily apparent when women resumed menses after postpartum amenorrhea because they visited a menstrual hut; they continued to do so periodically until they became pregnant. Thus, the waiting time to conception could be measured as the time from the first postpartum menstruation to the onset of pregnancy. The analysis of urinary steroid hormone profiles over a 10 week period showed that women in Sangui went to the menstrual huts during 87.5% of all menses and never went to the menstrual huts at other times (Strassmann, 1996). Compliance was nearly as good in the village of Dini. Interestingly, the Dogon themselves use knowledge of the timing of women's visits to the menstrual huts to monitor female reproductive status with a high degree of accuracy (Strassmann, 1992, 1996).

John et al. (1987: 437) conducted a particularly informative investigation of fecundability in Matlab Thana, Bangladesh. However, the timing of the first postpartum menses and the last menses before pregnancy were identified through interviews and therefore may have been subject to errors in recall or reporting. Although women were interviewed on a monthly schedule, "frequent absences of women on the interview day" resulted in "gaps of two to four months."

Other studies lacked information on menstrual cycling entirely and calculated fecundability as the time from marriage to 9 months before the first birth (e.g., Goldman et al., 1985). A disadvantage of this approach is that fecundability tends to be higher in the first than in subsequent conception waits on account of the relative youth of subjects, high coital frequency in the early years of marriage, and the absence of lactation (Goldman et al., 1985). For an unbiased measure of the waiting time to conception in the population as a whole, higher order conception waits need to included, which was possible in the present study.

METHODS

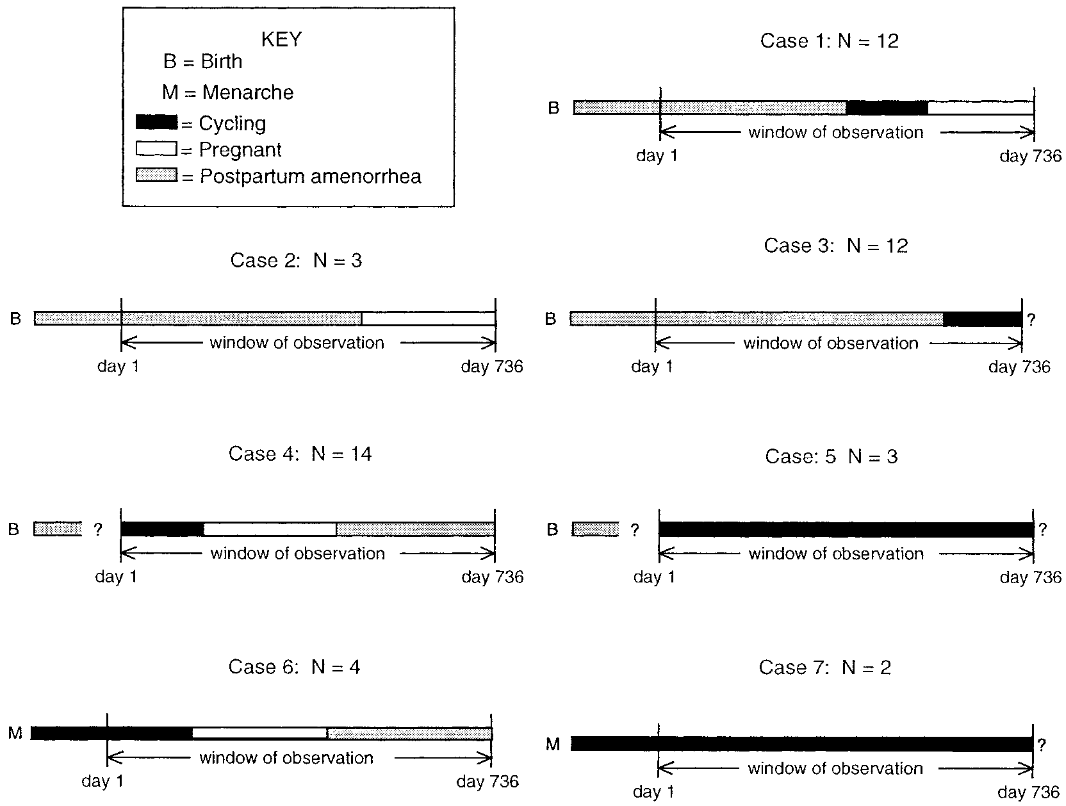
Research design

We chose a microdemographic approach emphasizing data quality rather than the more usual demographic survey approach in which large sample size is paramount. In particular, we focused on a single study village, Sangui, that had a population of 460 individuals in January 1988. Sangui is located on the plateau of the Bandiagara Escarpment at 14° 29′ N and 3° 19′ W. The data on conception waits are derived from a census of the women present at the two menstrual huts in the study village on each of 736 consecutive nights (July 1986 to July 1988). From the menstrual hut census we calculated the waiting times to conception for 50 women aged 16–41 years. Although this sample size is small, it is free of errors in recall or reporting that may arise from the use of retrospective data.

Because we wished to understand variation in conception waits in women actually exposed to the risk of conception, our sample of 50 women excludes the nonsusceptible women who cycled during the study. These nonsusceptible women (N = 12) were excluded for the following specific reasons: the husband had no children and was sterile (his wives (N = 2) subsequently divorced); the woman was sterile–younger than age 42 but no birth for past 5 years (N = 4); the husband was absent (N = 5); and illness (N = 1). Including these women in the analysis would have confounded the effort to determine how fecundability covaries with the independent variables of key interest. Eight women aged 42–53 years menstruated during the study but were probably sterile because, over an 8 year period (1986–1994), no woman in the study village had a live birth after age 41. We analyzed the data both with (N = 58) and without (N = 50) these eight women. The nonsusceptible women comprise 27% of the women who menstruated in Sangui at least once during the study, which may seem like a large proportion. Due to the absence of contraception, however, fecund women spent most (or all) of the 2 year study period pregnant or lactating, leaving subfecund and infertile women to be overrepresented among the cycling women (Strassmann, 1997). The sample of 50 women also excluded three women who could not be monitored: two were married to nonanimists and did not use the menstrual huts, and one was only intermittantly resident in the village.

The conception waits of only 15 of the 50 women were completely observed. For the remaining observations, the waiting times were censored (Fig. 1). Women of low fecundability were less likely to have conceived by the end of the study and therefore were likely to be overrepresented among the right censored observations. Because women of low fecundability cycle longer before pregnancy, they were also more likely to be cycling at the start of the study. As a consequence, they were also more prone to left censoring. Omission of either the right censored or the left censored observations would have resulted in an overestimate of fecundability. Three women conceived upon the resumption of ovulation postpartum without a prior menstruation. Exclusion of these women would have led to an underestimate of fecundability. The use of survival analysis made it possible to estimate fecundability from both the censored and uncensored waiting times (Lawless, 1982).

We used both continuous time and discrete time survival analysis. An advantage of the discrete time model is that it permits calculation of the probability of successful conception per menstrual cycle, but a disadvantage is that long and short cycles both count as one trial. The continuous time model has the advantage of being able to detect variation in conception waits caused by factors that impinge on cycle length, which should make this model more sensitive for assessing the effects of covariates. Our continuous time survival model is an exponential waiting time model where the hazard rate λ_i for woman i is given by $\lambda_i = \exp(-X_i\beta)$, where X_i is a vector of covariates (including a constant term) and β is a vector of regression coefficients. The exponential model was fit by maximum likelihood methods using Proc LIFEREG in SAS (1988). Proc LIFEREG handles right censored observations by standard methods (Kalbfleisch and Prentice, 1980; Lawless, 1982). We handled left censoring by truncating the waiting interval to the beginning of the study, which adjusted for length-biased sampling (Elandt-Johnson and Johnson, 1980; Wood et al., 1992; Tuma and Hannan, 1984).



Case	Likelihood for continuous time model		Likelihood for discrete time model	
1	$f_i(t)$	$t = t_f - t_o + 0.5$	$P_i(1 - P_i)^{n-1}$	$n = n_o$
2	$f_i(t)$	$t = 0.5$	$P_i(1 - P_i)^{n-1}$	$n = 1$
3	$s_i(t)$	$t = s_f - t_o$	$(1 - P_i)^n$	$n = n_o - 1$
4	$f_i(t)$	$t = t_f - s_o + 0.5$	$P_i(1 - P)^{n-1}$	$n = n_o$
5	$s_i(t)$	$t = s_f - s_o$	$(1 - P_i)^n$	$n = n_o - 1$
6	$f_i(t)$	$t = t_f - s_o + 0.5$	$P_i(1 - P_i)^{n-1}$	$n = n_o$
7	$s_i(t)$	$t = s_f - s_o$	$(1 - P_i)^n$	$n = n_o - 1$

Fig. 1. Types of censored and uncensored observations (N = 50) and corresponding likelihood functions. Continuous time model: t, waiting time (in months) to conception or censoring; t_o, time of the first visit to the menstrual hut during the study; t_f, time of the last visit to the menstrual hut during the study; s_o, time at which the study began; s_f, time at which the study ended. $f_i(t) = \lambda_i \exp(-\lambda_i t)$, $s_i(t) = \exp(-\lambda_i t)$, $\lambda_i = \exp(-X_i\beta)$ are the density function, the survival function, and the hazard rate of the waiting time to conception for the ith woman. Discrete time model: n, number of trials; n_o, number of observed visits to the menstrual hut. $P_i = \exp(X_i\beta)/1 + \exp(X_i\beta)$ is the probability of conception in a given cycle for the ith woman. X_i is the vector of covariates for the ith woman, and β is a vector of regression coefficients.

The discrete time survival model is a geometric model where the probability of conception per cycle, P_i, is given by $\exp(X_i\beta)/(1 + \exp(X_i\beta))$. This model was fit by standard logistic regression techniques (Proc LOGISTIC in SAS). Each menstrual cycle was considered one trial and was recorded as a success if it resulted in conception and a failure otherwise. Under the geometric model, the likelihood for a conception in cycle n is $P_i(1 - P_i)^{n-1}$. A right censored observation at n months will contribute likelihood $(1 - P_i)^n$. As with the continuous time model, left censoring of the conception risk interval was handled by left truncation to the beginning of the study period. The use of logistic regression in this context is described by Allison (1995).

In the continuous time model, we measured the waiting time to conception in units of 30 day months (rounded to the nearest tenth of a month). In the discrete time model, we counted the number of menstrual cycles that each woman experienced. (The correlation coefficient between the waiting time to conception in months and the number of menstrual cycles was 0.93, $P < 0.0001$, $N = 50$ women). Under the continuous time model, left censored observations were measured from the beginning of the study, and right censored observations were measured up to the end of the study. Under the discrete time model, left censored observations were measured from the first menstruation during the study, and right censored observations were measured up to the last menstruation during the study.

The definition of the end of postpartum amenorrhea required special consideration. In natural fertility populations, whether menstruation usually precedes the first postpartum ovulation depends on the duration and intensity of breast-feeding (Perez et al., 1972). When amenorrhea lasts a median of 20 months and supplemental foods are not introduced until a child is 1 year of age, as in the Dogon case, then it is more common for menstruation to precede ovulation (see Udesky, 1950; Tietze, 1963; Perez et al., 1972). Even when the first ovulation occurs prior to the first menstruation, endocrine abnormalities usually preclude the establishment of a viable pregnancy before the second cycle (Diaz et al., 1992). Three women in our study did, however, conceive before they made a single postpartum visit to the menstrual hut. The Dogon are aware of the rarity of such conceptions and assign a special name to a child conceived so promptly (*Akunyon* for a boy and *Yakunyon* for a girl). We assumed that postpartum amenorrhea ended with the first postpartum menstruation, except for the three observations where conception actually preceded menstruation. It would have been useful to estimate the diminished fecundability of the month previous to the first postpartum menstruation, and we attempted to do so, but our sample size was inadequate for this task. The cycle in which conception occurred also required special attention. In the continuous time model, we counted the time from the last menstruation to conception as 0.5 months. In the discrete time model, the same period corresponds to an additional trial. Figure 1 shows how these definitions and decision rules apply to our data.

A covariate was included in the final continuous time model if it was significant according to the likelihood ratio test ($\alpha = 0.05$). To permit comparison, the final discrete time model included the same covariates as the continuous time model. The small sample properties of maximum likelihood estimation may be in doubt; therefore, we used a Gibbs sampler to obtain exact small sample inferences (Gilks et al., 1996). The posterior means and standard deviations obtained through the Gibbs sampler agreed very closely with the maximum likelihood parameter estimates and their standard errors.

For the continuous time model, we assessed the importance of the covariate effects by calculating the relative risk of conception, e^{-b}, for a one unit increase in each covariate (b is the coefficient of the covariate under consideration). For the discrete time model, we calculated the odds ratio for conception, e^{b}. We also calculated fecundability, f, which can be expressed in terms of the exponential hazard rate, λ_i, as $f = 1 - \exp(-\lambda_i)$. For the discrete time model, fecundability is directly available as P_i. Finally, we plotted fecundability against wife's age using age-adjusted means for each of the significant covariates in the continuous time model.

Independent variables

The independent variables that were tested for a significant association with the waiting time to conception fall into five categories: 1) life history covariates, 2) lactation, 3) nutrition, 4) economics, and 5) polygyny and marital status.

Life history covariates. These include age of both spouses, marital duration, gravidity (number of previous pregnancies), and parity (number of previous births). Gravidity was a continuous variable (range 0–10) representing the number of pregnancies a woman had experienced prior to the study. Parity was a binary variable (0 = nulliparous, 1 = parous) that enabled us to compare the results for the six nulliparous women in the study with those for the 44 parous women. The above information was determined from interviews using a questionnaire on marital and reproductive histories. The questionnaires were pretested on a subsample and cross-checked extensively. The age estimates are highly accurate because Dogon males and females belong to narrow age classes of two and a half or about 2.5 years (one birth interval); they even keep track of birth order within each age class. At the time of the interviews, B.I.S. had lived more than 2 years in the study population, spoke fluent Dogon, and knew each of the respondents individually. This familiarity enhanced the goodwill of the subjects and their willingness to respond accurately. Indirect remuneration in the form of first aid and assistance with a "self-help" development project also promoted subject cooperation (for details see Strassmann, 1996).

Lactation. Three covariates provided potential measures of breast-feeding: 1) nursing status, 2) duration of postpartum amenorrhea, and 3) sex of the last child. Nursing status is a categorical variable because it was recorded as present or absent, but its intensity was not measured. The duration of postpartum amenorrhea is a proxy for the intensity of nursing because it equals the age of the nursing child upon the resumption of menstruation. Among the Dogon, older children consume more supplemental food and nurse less. Sex of the last child is expected to be a significant covariate of conception waits if there is a sex difference in nursing.

Nutrition. The nutritional covariates include 1) the body mass index (wt/ht^2), 2) the sum of the triceps, subscapular, and lateral calf skinfold thicknesses, and 3) a binary variable that distinguished the women who were undernourished from those who were in the normal range. This variable was based on the recommendations of Frisancho (1990), who defined women aged 18–45 years as undernourished if the sum of the triceps and subscapular skinfolds was 20.0 mm or less. The women were measured during the hunger season in June 1987 and July 1988 when most families had already consumed the millet and other cereal crops from their previous harvests and were living off purchased grain. Women supplemented their diets with wild berries or collected firewood to sell at the market in exchange for small bowls of rice. Thus, at the time of the nutritional status measurements, the population appeared to be chronically food-stressed. The women were measured at approximately the same time of year, which controlled for seasonal variation in nutritional status. All measurements were made by a single individual, preventing interoperator error. To control for reproductive status, we excluded all the women who were not measured during cycling.

Economics. In a population of subsistence farmers subject to food shortages, a woman's nutritional status may depend on her economic status. We therefore quantified the resources of each of the 59 different economic units in the village. We call these economic units work-eat groups because they are composed of the people (men, women, and children) who cultivated the same millet fields and who assembled in one compound to eat together. The economic status of a Dogon woman depends on that of her work-eat group. A married woman always belongs to the work-eat group of her husband, and she does not have a separate avenue to wealth. Work-eat groups are not comparable to families because people have close relatives in other work-eat groups. Nor are they comparable to households because members of the same work-eat group often sleep in different compounds. For example, adolescents often do not sleep in the same compound as their parents, and cowives may have different compounds. Married men who are close relatives may belong to the same work-eat group but do not share the same compound. The significance of a work-eat group lies in the economic interdependence of its members, who share the same standard of living.

The Dogon are subsistence farmers, so economic status depends on ownership of food resources: millet and other cereal crops and to a lesser degree onions and livestock. We measured the 540 cereal fields (millet, sorghum, rice, fonio) and 422 onion fields belonging to the 59 work-eat groups. We calculated field areas from angles and sides using a trigonometric calculator program. To calculate yields, we assigned a value of 1 to 5 to each grain field to indicate the relative quality of the standing crop just before harvest; then we weighed the grain in representative 25 m^2 plots. Foot traffic to and from the fields must funnel across a bridge. An observer posted at this bridge throughout the 1987 harvest tallied all baskets of grain carried. We weighed baskets representative of each size class. This procedure gave a second and independent assessment of each work-eat group's yield in kilograms. The primary livestock of the Dogon are sheep and goats, but the wealthiest families can afford cattle. Other animals include donkeys, horses, and pigs. We counted each of these animals by sex and age (juvenile vs. adult).

About 5% of the economy of Sangui is composed of income from commerce (see below). Some men buy onions and resell them to middlemen who sell them in the West African cities of Bamako and Abidjan. Two men went to Bamako themselves to sell onions. One man regularly biked the 120 km to Mopti on the Niger River to buy dried fish to resell at the local market. Another man was a butcher. A Dogon assistant who is familiar with local commerce helped us to estimate the profit for each man.

To calculate the total economic assets of each work-eat group, we estimated the monetary value of each group's cereals, onions, livestock, and revenues from commerce. These estimates are based on the prevailing local market prices for the various commodities at the time of the fieldwork. Monetary values are expressed in CFA, which is a West African currency tied to the French franc. During the study period, 50 CFA equaled 1 FF. We then totaled the assets of the 59 work-eat groups to estimate the resource holdings for the entire village.

We computed seven different economic indices for each work-eat group because it was unclear which would be the most informative. With the exception of the third and seventh index, we divided the resources of each group by the estimated daily energy requirement of its members in megajoules (MJ). These requirements depend on the age and sex composition of each group and are based on the recommendations of the FAO/WHO (1973). The economic indices are as follows:

ECON1 = cereal hectarage/MJ;

ECON2 = cereal CFA/MJ;

ECON3 = cereal CFA/cereal hectarage;

ECON4 = (cereal + onions + livestock + commerce) CFA/MJ;

ECON5 = [cereal + onions + 0.07(livestock) + commerce] CFA/MJ;

ECON6 = [cereal + onions + 0.07(livestock) + 0.07(commerce)] CFA/MJ;

ECON7 = Economic rank of the work-eat group relative to the others.

The first index is the hectarage of cereal crops planted divided by the energy requirements. The second is the value of the cereal yield in CFA divided by the energy requirements. The third index is the CFA value of the cereal yield per hectare, which is a measure of land quality. The fourth index totals the CFA value of all resources, giving equal weight to each type, and divides the sum by the energy requirements. The fifth index is the same, except it weights the value of livestock by a factor of 0.07 to reflect only the estimated value of milk and meat consumed or sold. The sixth index is the same as the fifth, except it also weights the revenues from commerce by a factor of 0.07 since most of the commercial revenues do not enter the family food budget. Finally, the seventh index is the average rank of each work-eat group with respect to the others across the following five variables: cereal hectarage, onion hectarage, value of cereal yield, value of livestock, and revenues from commerce. The wealthiest group had a rank of 59, and the poorest group had a rank of one.

Polygyny and marital status. We expressed the degree of polygyny in each work-eat group as the number of married males divided by the number of married females. This index ranged from 0.33 (three wives per man) to 1.0 (one wife per man). Another variable categorized a woman's marital status as monogamously or polygynously married. A binary variable distinguished the married women from the fiancées to permit us to test the assumption that the fiancées were fully exposed to the risk of conception. A distinct marriage ceremony is absent, and it is normative for regular sexual relations with a fiancé (and preferably the birth of one to three infants) to precede cohabitation.

Educational status. Only two of the adult men and none of the women in the village could read and write. Since very few individuals had completed even one year of primary school education, we did not include educational status as an independent variable.

RESULTS AND DISCUSSION

Thirty-three of the 50 women became recognizably pregnant during this study, and 25 women cycled without becoming pregnant (Fig. 1). Twenty-five of the women who became recognizably pregnant actually delivered during the study, and eight delivered after the 736 day period of prospective observation was over. Confirmation of the latter deliveries (seven were live births, and one was a stillbirth) took place through interviews on the next field trip. Short pregnancies that ended in miscarriage during the first trimester were not detected because, in the menstrual hut census, such pregnancies could not be distinguished from cycle irregularity. Fecundability as measured in this study is best described as effective fecundability, although one infant was stillborn (see Wood 1994).

Table 1 lists the mean, standard deviation, and number of women for whom data were available for a particular variable. Table 2 shows the final continuous and discrete time models. The coefficients in the two models are very similar with respect to magnitude but are opposite in sign. This occurs because the discrete time model estimates the cycle-wise odds of conception (O_i) for the ith woman as exp ($X_i\beta$), whereas the continuous time analysis models the hazard rate (λ_i) for the ith woman as exp ($-X\beta$). When the cycle-wise probability of conception is small, O_i and λ_i are approximate reciprocals.

The continuous time model produced more highly significant results than the discrete time model for six of the seven covariates (Table 2). This was expected because the continuous time model takes variability in cycle length into account, whereas the discrete time model does not. Instead, each cycle in the discrete model was counted as one trial, no matter how long or short. Natural variation in cycle length for a given woman is associated with anovulation. Anovulatory episodes were particularly common among the perimenopausal women and the youngest women. Occult pregnancies that went unrecognized were probably another major cause of variability in cycle length, especially among the most fecund women. Hormonal data indicated that, in the village of Sangui, women did not go to the menstrual hut during 12.5% of all menses, though they never were observed to miss two visits in a row (Strassmann, 1996). Missed visits introduce more noise in the discrete, cycle-wise analysis than in the continuous time analysis. Missed visits should not affect the continuous time analysis unless they occur at the beginning or end of the conception wait. If missed visits were frequent at such times, then one would expect that some of the women who were ostensibly in postpartum amenorrhea during the entire study and not using the huts were in fact cycling. However, the hormonal data disclosed no such cases (Strassmann, 1996). In sum, natural variation in cycle length and the difficulty of detecting 100% of all menses over a 2 year period probably explain why the continuous time model gave a better fit—as assessed by smaller P values and smaller standard errors for the significant covariates.

We also fit a discrete time model in which the number of trials for each woman was

TABLE 1. Independent variables tested for an association with the waiting time to conception

Variable	N (women)	Mean	SD
Wife's age (years)	50	27.72	6.80
Husband's age (1 = <35, 2 = 35–49, 3 = >49 years)	50	36.06[1]	9.40[1]
Marital duration (years)	50	4.38	5.66
Gravidity (number of prior pregnancies)	50	3.3	2.57
Parity (0 = nulliparous, 1 = parous)	50	0.88	0.33
Nursing (0 = no, 1 = yes)	50	0.58	0.50
Duration of postpartum amenorrhea (months)	20	20.10	7.22
Sex of last child (0 = male, 1 = female)	44	0.34	0.48
Nutritional status			
Weight (kg)/height-squared (m)	24	21.18	1.43
Sum of triceps, subscapular, and lateral calf skinfolds (mm)	24	24.63	5.89
Sum of triceps + subscapular skinfolds (0 = 20.00 mm or less, 1 = greater than 20.00 mm)	24	0.71	0.46
Economic status[2]			
ECON1	50	278.64	81.34
ECON2	50	751.46	273.84
ECON3	50	0.04	0.02
ECON4	50	3,349.66	1,866.67
ECON5	50	2,140.46	788.65
ECON6	50	1,943.52	600.63
ECON7	50	40.42	12.05
Polygyny (number of married males/married females)	50	0.60	0.19
Cowives (0 = one or two, 1 = none)	44	.16	0.37
Marital status (0 = fiancée, 1 = married)	50	0.76	0.43

[1] Mean and standard deviation for the continuous variable, the husband's age in years.
[2] See text for definition of the seven economic indexes.

calculated from her waiting time to conception in days (in the continuous time model) divided by 30 days. The results were more similar to those from the continuous time analysis than to those from the discrete time analysis in Table 2. We now turn to a discussion of the covariates that were significant in the continuous time model at $\alpha = 0.05$, including wife's age, wife's age squared, husband's age, marital duration, gravidity, and nursing.

Significant covariates

Wife's age. The negative coefficient for age and the positive coefficient for age-squared in the continuous time model (and the reverse in the discrete time model) indicate that conception waits had a U-shaped relationship with age. Consistent with previous studies (Bendel and Hua, 1978; Jain, 1969a; Wood et al., 1994), fecundability had an inverse U-shaped relationship with age. In Figure 2, fecundability is plotted against wife's age using age-adjusted means for the other significant covariates. Age-adjusted means for gravidity and marital duration were obtained from linear regressions of these variables on age. Age-adjusted means for the ordinal variable, husband's age, were obtained from an ordinal logistic regression. (Age-squared was not significant in any of these regressions.) A logistic regression of nursing status on age was not significant, so nursing status was estimated by its population mean. Figure 2 also shows the upper and lower 95% confidence limits for the age-specific values of fecundability. These limits were obtained from the estimated covariance matrix of the parameter estimates produced by the LIFEREG procedure in SAS. The age-specific changes in fecundability shown in Figure 2 may reflect changes in coital frequency, fecundity, and fetal loss.

As shown in Table 3, the age-specific values for effective fecundability were similar in the continuous time model whether or not the eight women older than 41 years were included or excluded from the sample. The estimated peak fecundability occurred from age 26–29 and had the following values: 0.18 (N = 58) and 0.19 (N = 50). In the discrete model, peak fecundability was 0.13 (N = 58) and 0.21 (N = 50). When all covariates were at their population means, the monthly probability of conception for the 50 women in the continuous time model was 0.11, which corresponds to a waiting time to conception of

TABLE 2. Independent variables significantly associated with the waiting time to conception

Independent variables	Continuous time model (N = 58)[1]		Continuous time model (N = 50)		Discrete time model (N = 58)[1]		Discrete time model (N = 50)	
	Coefficient	*P*	Coefficient	*P*	Coefficient	*P*	Coefficient	*P*
Wife's age (years)	−0.601	0.006	−0.553	0.019	0.485	0.055	0.446	0.101
Wife's age-squared (years)	0.011	0.003	0.010	0.011	−0.009	0.027	−0.009	0.056
Husband's age (<35 years)[2]	−2.035	0.005	−1.982	0.005	2.087	0.008	2.058	0.008
Husband's age (35–49 years)[2]	−1.477	0.029	−1.398	0.037	1.795	0.011	1.753	0.012
Marital duration (years)	0.183	0.002	0.179	0.003	−0.182	0.028	−0.181	0.032
Gravidity (number of prior pregnancies)	−0.550	0.010	−0.543	0.011	0.551	0.056	0.551	0.060
Nursing (0 = no, 1 = yes)	−0.741	0.052	−0.744	0.051	0.653	0.126	0.655	0.125
Intercept	12.652	—	11.984	—	−10.809	—	−10.268	—
−2 log likelihood	129.094	—	128.518	—	179.656	—	179.317	—

[1] This model includes eight women, aged 42–53.
[2] Compared to the omitted category, husband's age >49 years.

8.3 months. In the discrete time model (N = 50), the probability of conception per menstrual cycle was 0.14, which corresponds to an expected waiting time of 7.3 months.

These results can be compared with those for Western women who have undergone artificial insemination. The inseminations take place near the time of ovulation, which controls for any effect on fecundability that is due to coital frequency. In the Netherlands, the cumulative pregnancy rate after 12 menstrual cycles of insemination was 0.75 (for women aged <24 years), 0.72 (25–29 years), 0.72 (30–34 years), and 0.49 (>34 years) (van Noord-Zaadstra et al., 1991). These values correspond to fecundabilities of 0.11 (<24 years), 0.10 (25–34 years), and 0.06 (>34 years). The lower fecundability of the Dutch women may be due to a number of factors, including the use of frozen semen and a tendency for less fecund women to use artificial insemination.

John et al. (1987) found that Bangladeshi women who were fully breast-feeding their infants upon the resumption of menses had a monthly probability of conception of about 0.01. Women who had begun to wean their infants 4–6 months prior to the resumption of menses had a probability of conception of 0.06, and women who had ceased breast-feeding just before the resumption of menses had a monthly probability of conception of 0.19. Other studies report higher fecundability, ranging from 0.14–0.31 (Wood, 1989). However, they measure the time from marriage to 9 months before the first birth and therefore include only young, nonlactating women with short marital durations, resulting in a bias toward high fecundability.

Husband's age. Figure 3A plots fecundability against wife's age for three different values of the husband's age, with marital duration and gravidity at their age-specific means and nursing at its population mean (nursing status did not vary with age). According to the coefficients for husband's age in the continuous time model (Table 2) (N = 50), if the husband was younger than 35 years, the relative risk of conception was 7.3 ($P = 0.005$) compared with couples in which the husband was 50 years or older, after controlling for the other significant covariates. When the husband was 35–49 years old, the relative risk of conception was 4.0 ($P = 0.04$) compared with couples in which the husband was 50 years or older (Fig. 3A). (The corresponding odds ratios in the discrete time model, N = 50, were 7.8 and 5.8.) Most investigators disregard the effect of husband's age on fecundability on the assumption that it is small in relation to the effect of wife's age and other covariates (see Goldman and Montgomery, 1989). Our results demonstrate that husband's age is

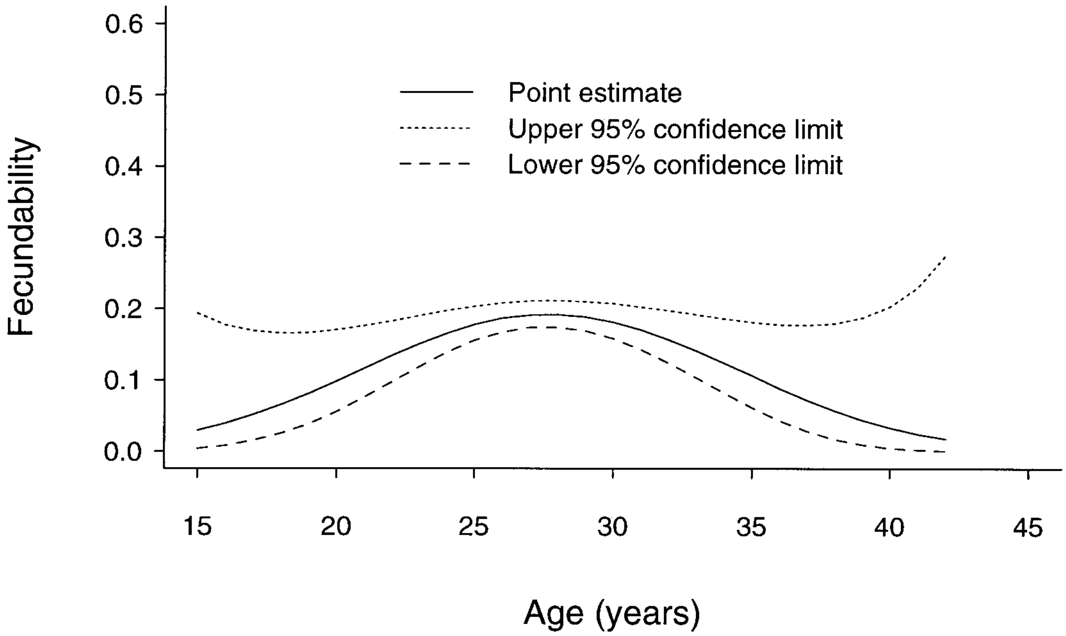


Fig. 2. Fecundability by wife's age, with the other significant covariates at their age-adjusted means (continuous time model, N = 50 women). The broad confidence intervals at the extremes of the age distribution are due to small sample size (below age 20, N = 5; above age 40, N = 2). When the eight women older than 41 years are included, the confidence intervals become tighter between ages 30 and 40 years and wider between ages 15 and 25 years.

TABLE 3. Effective fecundability by age

	Continuous time model		Discrete time model	
Age	N = 58	N = 50	N = 58	N = 50
15	0.03	0.03	0.03	0.05
20	0.10	0.10	0.08	0.13
25	0.17	0.18	0.12	0.20
30	0.17	0.18	0.12	0.20
35	0.08	0.11	0.07	0.13
40	0.02	0.03	0.02	0.05

an important predictor of fecundability, regardless of whether the sample included women aged 16–41 years (whose husbands ranged in age from 21–57 years) or women aged 16–53 years (whose husbands were 21–64 years). The relative contributions of coital frequency, deteriorating health, and reproductive senescence, all of which are interrelated, remain a problem for the future. A selection bias may also exist such that less fecund women are more likely to marry men who are much older. In this study, however, the potential for such a bias was lessened by the exclusion of women who were prematurely sterile (younger than age 42 but had not given birth in the past 5 years). Although we cannot identify the mechanism behind the effect for husband's age, these results empirically support the suggestion (Campbell and Leslie, 1995) that human reproductive ecologists need to consider the male contribution to reproduction.

Marital duration. The waiting time to conception increased with marital duration. This effect is shown in Figure 3B, in which fecundability is plotted against wife's age with marital duration, husband's age, and gravidity at their age-specific means and nursing at its population mean. Figure 3B also shows the relationship with fecundability when marital duration is half a standard deviation above or below its age-adjusted mean. Based on the coefficient for marital duration in the continuous time model (Table 2) (N = 50), a 1 year increase in marital duration was associated with a relative risk of conception of 0.84 after controlling for wife's age, husband's age, gravidity, and

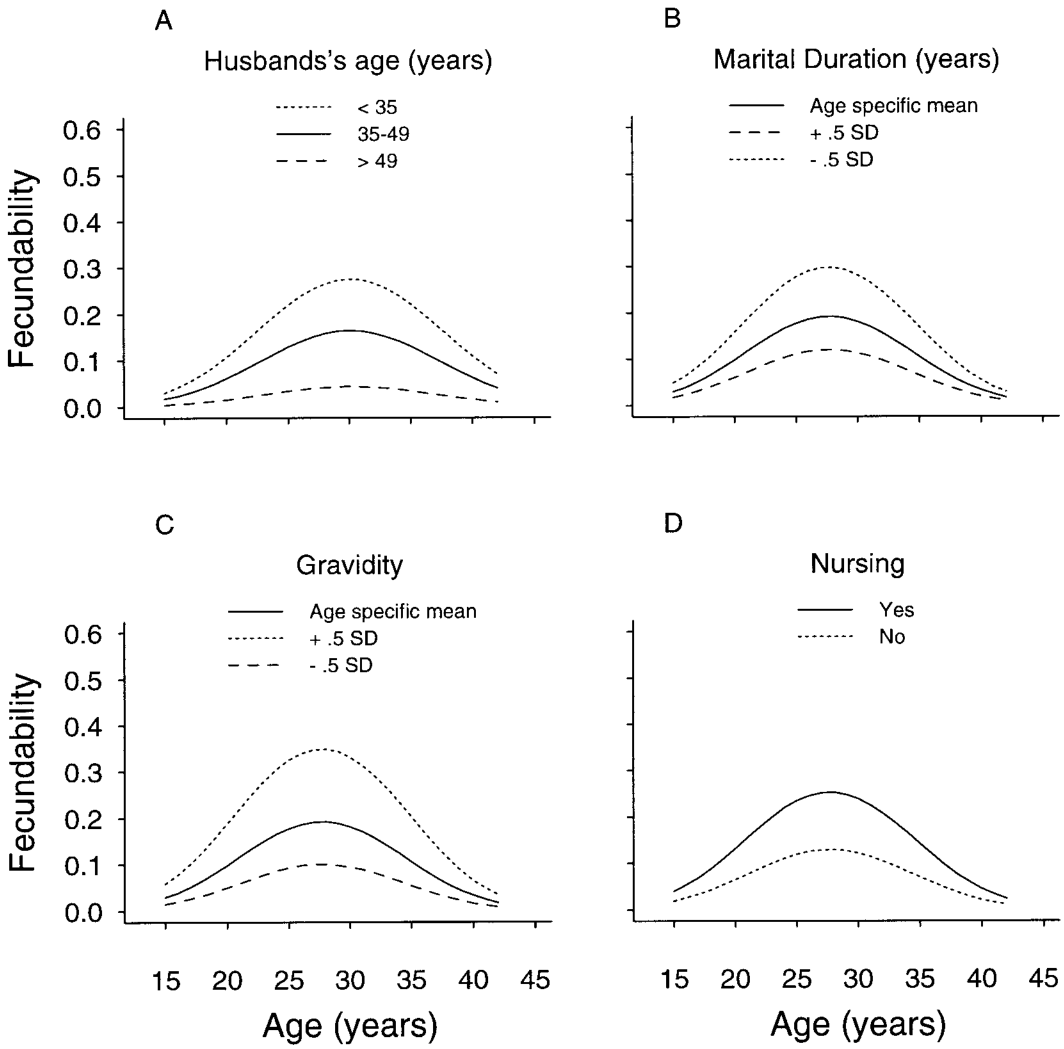


Fig. 3. Fecundability by wife's age plotted against (**A**) husband's age (<35 years, 35–39 years, and 40+ years), (**B**) age-adjusted mean marital duration ± 0.5 standard deviation (SD), (**C**) age-adjusted mean gravidity ± 0.5 SD, and (**D**) nursing status (yes or no). In each graph, the other significant covariates are at their age-specific means with the exception of nursing, which did vary with age. Continuous time model, N = 50.

nursing. (The corresponding odds ratio in the discrete time model was 0.83.) Since women had a median of two husbands in a lifetime (Strassmann, 1992), many of the older women in the sample had been married for fewer years than the younger women. Thus, the significant coefficient for marital duration cannot be considered simply an age effect. The frequency of coitus tends to decrease rapidly during the first few years of marriage followed by a more gradual decrease in later years (James, 1983), suggesting a likely mechanism for the relationship between fecundability and marital duration observed in this study.

Previous analyses of fecundability using multivariate survival analysis have seldom included marital duration as an independent variable. However, simpler analyses have found a linear increase in conception waits with marital duration. One such study used World Fertility Survey data from Africa, Latin America, Southern Asia, and the Middle East (Goldman et al., 1987).

Gravidity. When gravidity was the sole covariate, its coefficient was positive in the continuous time model due to confounding with other covariates. After controlling for age of the spouses, marital duration, and nursing status, the coefficient for gravidity became negative, indicating that women who had many previous pregnancies had shorter conception waits. The relationship between fecundability and the wife's age is shown in Figure 3C for three values of gravidity: the age-specific mean and half a standard deviation above and below the age-specific mean. Based on the coefficient for gravidity in the continuous time model (Table 2) ($N = 50$), the relative risk of conception for each additional prior pregnancy was 1.7. (The corresponding odds ratio in the discrete time model was also 1.7.) Nonetheless, it is unlikely that higher gravidity causes shorter conception waits. Instead, gravidity is probably correlated with additional, unmeasured variables that have a direct influence on fecundability. Although the association between gravidity and fecundability is unlikely to be causal, it is clear that in the study village a woman's past fecundability is a good predictor of her future fecundability. This association has been widely recognized in preindustrial societies and explains why people judge a woman's fecundity from her past reproductive performance.

If maternal somatic reserves become depleted over a succession of pregnancies, the so-called maternal depletion syndrome (Jelliffe and Maddocks, 1964; Tracer, 1991), then a reduction in fecundability might be expected. The magnitude of the relationship should depend on the sensitivity of fecundability to nutritional status or other measures of maternal depletion. Our results do not refute the concept of maternal depletion but suggest that, if this phenomenon is occurring in the study population, then either it does not affect the waiting time to conception or the effect is so small it is masked by other factors. Rather than a decline in fecundability with increasing gravidity, we see an increase in fecundability.

Popkin et al. (1993) found that higher parity predicts longer conception waits. However, Popkin et al. may not have entirely controlled for the confounding variable, maternal age, because they treated age as a dichotomous rather than a continuous variable. They also did not control for marital duration. Adequate controls for both age and marital duration are essential when testing for an association between parity (or gravidity) and conception waits. When we omitted these two controls from our models, we obtained the misleading result that women of higher gravidity had longer conception waits. Perhaps the coefficient for parity in Popkin et al.'s study would also change sign if they controlled for marital duration and specified age as a continuous variable. Larsen and Vaupel (1993) reported that fecundability in North American Hutterites declined by about 25% from parity zero to parity one and then approximately leveled off at higher parities. However, marital duration was not controlled, and the standard errors for the estimates were large.

Nursing. Nursing status was a dichotomous variable, with 0 indicating that a woman was not nursing and 1 indicating that she was nursing. In the continuous time model (Table 2) ($N = 50$), the relative risk of conception was higher by a factor of 2.1 if a woman was nursing after controlling for age of both spouses, marital duration, and gravidity. (In the discrete time model, the corresponding odds ratio was 1.9.) Figure 3D plots fecundability by wife's age and nursing status with the other significant covariates at their age-specific means. The finding that fecundability was higher in nursing women is surprising and at first appears contradictory to the results of previous studies (e.g., John et al., 1987; Goldman et al., 1987; McNeilly et al., 1982; Singh et al., 1993). However, a closer examination of the data offers an explanation. Dogon women nurse throughout the entire waiting time to conception and sometimes even nurse throughout the subsequent pregnancy. The consequence is that any woman in our study who was not nursing was either 1) less at risk for conception or 2) her previous infant had died. Thus, when a sample is split into the women who are nursing and the women who are not nursing, it is also thereby dichotomized into a more at risk group and a less at risk group. Lactation was also associated with shorter conception waits among Cebu women of the Philippines (Popkin et al., 1993); we suggest that the same explanation may apply.

Nonsignificant covariates

To verify that age was the only covariate with a significant quadratic effect, we also tested marital duration–squared and gravidity-squared in the final model, but neither was significant. Interactions between age and marital duration, gravidity, and nursing also were not significant. Other independent variables that were not significantly associated with conception waits are presented in Table 4. Each of these covariates was individually adjusted for the significant covariates in Table 2. We will discuss each of the nonsignificant variables in turn.

Duration of postpartum amenorrhea. The duration of postpartum amenorrhea was not significantly associated with the waiting time to conception. However, after we excluded cases in which the infant died or the duration of postpartum amenorrhea was censored, the sample size was reduced to 20 women. The use of techniques for the analysis of data in which both the dependent and the independent variables are censored would have seriously complicated the analysis. In previous studies, longer periods of postpartum amenorrhea were associated with an increased monthly probability of conception (John et al., 1987). However, this relationship was not found among women who breast-fed for at least 9 months after the resumption of menses (Singh et al., 1993).

Sex of the last child. The coefficient for sex of the last child implies that the birth of a daughter (instead of a son) is associated with a relative risk of conception of 1.7 (continuous time model) and 2.1 (discrete time model) after controlling for the significant covariates. However, the coefficient was not significant in either model. Conspicuous sex-biased investment in infants, such as female infanticide, was absent in the study village. We cannot exclude the possibility of more subtle forms of sex bias in some families. For example, perhaps daughters are sometimes nursed less intensively than sons, or they may of their own accord suckle less.

Nutritional status. After we excluded individuals who were not measured during cycling, we had cross-sectional anthropometric data on only 24 women, which is a small sample. Nonetheless, this sample includes almost half of the fecundable women in the study village over a 2 year period. Therefore, it is reasonably large with respect to the population to which we are drawing inferences. Seven of the 24 women met the criterion for undernourishment (combined thickness of the triceps and subscapular skinfolds <20 mm). This variable ($0 =$ normal, $1 =$ undernourished), the body mass index (wt/ht^2), and the sum of three skinfolds (triceps, subscapular, and lateral calf in millimeters) were not significant in either model.

Similarly, in Bangladesh and Guatemala, no significant correlation was found between the mean waiting time to conception and anthropometric measures (Chowdhury, 1978; Bongaarts and Delgado, 1979). Caloric intake, measured during a nutritional intervention program in Guatemala, also did not affect mean waiting times to conception (Delgado et al., 1982). In rural Bangladesh, neither height nor body mass index (wt/ht^2) was related to fecundability (John et al., 1987). A significant association was found between nutritional status and both the risk of intrauterine mortality and the duration of postpartum amenorrhea, but no association was found between nutritional status and the probability of conception during cycling (Ford et al., 1989).

Extreme energy deprivation, such as occurs during famines, has a demonstrable effect on fecundity (Bongaarts, 1980; Stein and Susser, 1978). Moreover, clinical and endocrinological studies demonstrate that ovarian function is constrained by energy balance (Warren, 1983; Ellison et al., 1993). Nonetheless, subtle forms of undernutrition do not appear to be a major cause of variation in conception waits within populations.

TABLE 4. Nonsignificant covariates individually adjusted for the significant covariates in Table 2 (N = 50)[1]

Independent variables	Continuous time Coefficient (standard error)	Continuous time P	Discrete time Coefficient (standard error)	Discrete time P
Parity (0 = nulliparous, 1 = parous)	0.765 (1.033)	0.459	−1.384 (1.153)	0.230
Duration of postpartum amenorrhea (months)	0.018 (0.076)	0.815	−0.008 (0.077)	0.916
Sex of last child (1 = male, 2 = female)	−0.553 (0.461)	0.231	0.737 (0.521)	0.187
Nutritional status				
Body mass index (wt/ht^2)	0.096 (0.269)	0.722	−0.279 (0.300)	0.353
Sum of three skinfolds (mm)	0.039 (0.079)	0.618	−0.166 (0.109)	0.129
Normal = 0, undernourished = 1	0.361 (0.914)	0.693	−0.849 (0.985)	0.389
Economic status				
ECON1	0.009 (0.023)[2]	0.686	−0.015 (0.028)[2]	0.579
ECON2	0.152 (0.076)[2]	0.045	−0.081 (0.087)[2]	0.349
ECON3	0.015 (0.030)[2]	0.616	−0.012 (0.036)[2]	0.746
ECON4	0.008 (0.010)[2]	0.416	−0.008 (0.012)[2]	0.501
ECON5	−0.003 (0.241)[2]	0.990	0.140 (0.308)[2]	0.650
ECON6	19.300 (12.416)	0.120	−10.403 (13.258)	0.433
ECON7	0.015 (0.014)	0.288	−0.017 (0.017)	0.320
Polygyny index (number of married males/ number of married females)	−0.566 (1.058)	0.593	−0.052 (1.278)	0.968
Cowives (0 = one or two, 1 = none)	0.726 (0.537)	0.176	−1.200 (0.662)	0.070
Marital status (0 = fiancée, 1 = married)	1.083 (0.714)	0.129	−0.971 (0.819)	0.236

[1] Coefficients and standard errors are for an increase of one unit unless otherwise indicated.
[2] For an increase of 100 units.

Economic status. The total economic resources of the study village are shown in Table 5. Accurate and objective measures of economic status were a priority of the present study, but among the seven economic indexes only ECON2 (CFA value of the cereal crop/MJ) was significantly associated with the waiting time to conception. According to ECON2, women in wealthier work-eat groups had longer conception waits ($P = 0.045$) in the continuous time model, but no significant effect was found for ECON2 in the discrete time model. Given that six of the seven indices were not significant, economic status does not appear to be a good predictor of fecundability in our study population. The heterogeneity of the sample was substantial (Table 1), with some women belonging to work-eat groups that had surplus wealth in the form of cattle, and others belonging to work-eat groups in which hunger was a significant threat. Women in poorer work-eat groups tended to work harder (Strassmann, unpublished data), but we detected no effect on conception waits. Economic differences might be expected to influence fecundability through effects on nutritional status, coital frequency, or breastfeeding, but we found no evidence for these effects.

Jain (1969b) reported that fecundability among Taiwanese women was positively associated with socioeconomic status after controlling for age and marital duration in a multiple regression. In a hazards analysis, John et al. (1987) found no relationship between economic status, as determined from interviews, and fecundability in rural Bangladesh. Ford et al. (1989) found that two indicators of socioeconomic status, religion and husband's occupation, were not associated with the waiting time conception. In Uttar Pradesh, India, an index of household social status was not significantly related to conception waits in multivariate survival models (Singh et al., 1993; Nath et al., 1994). Upper-caste Hindus, however, had longer conception waits than lower-caste Hindus and Muslims. With the exception of this last result, survival analyses have found that economic status is not an important determinant of the variance in conception waits within populations.

Marital status and polygyny. Fiancés had shorter conception waits than married women (although not significantly so), indicating that it was appropriate to include the fiancés in the at risk group. The maximum number of wives for a given man was three, and the mean ± standard deviation for the ratio of married males to married females in a work-eat group was 0.60 ± 0.19. This ratio was not a significant predictor of the waiting

TABLE 5. Economic resources of Sangui in 1987

Resource	Quantity	CFA[1] value
Cereal crops		
Grain	42,000 kg	3,150,000
Land	127 ha	—
Yield	330 kg/ha	—
Onions		
Bulbs and stems	—	4,160,000
Land	4 ha	—
Livestock		
Cattle	89 head	—
Sheep	141 head	—
Goats	269 head	—
Other	12 head	—
Total	511 head	6,070,000
Commerce	—	740,000
Total	—	14,120,000

[1] Approximately 310 CFA/dollar in 1987.

time to conception. Similarly, the number of cowives (0 = one or two, 1 = none) did not predict a woman's waiting time, although this variable approached significance in the discrete model.

Demographic studies have sometimes (Dorjahn, 1958; Bean and Mineau, 1986) but not always (Chojackna, 1980; Borgerhoff Mulder, 1989) found that polygyny is associated with lower female fertility. Before the possibility that polygyny causes reductions in female fertility can be resolved for any particular population, two confounding issues need to be addressed: 1) marriages of low fertility may be more likely to become polygynous, and 2) women of low prior fertility may be more likely to become junior wives in polygynous unions (Pebley and Mbugua, 1989). It is difficult to control for these possibilities using retrospective data on fertility. However, the absence of a significant relationship between polygyny and conception waits in our study, in which conception waits were monitored prospectively, implies that such a relationship did not exist in our study population.

CONCLUSIONS

This study used survival analysis to identify the significant covariates of the waiting time to conception in a Dogon village in Mali, West Africa. Through microdemographic research on a small but high-quality sample, we gained five distinct methodological advantages: 1) the calculation of conception waits from prospective data on menstrual cycling, 2) the inclusion of conception waits in both nulliparous and parous women as well as conception waits in women who went directly from amenorrhea to pregnancy, 3) unusually accurate data on marital and reproductive histories obtained by the anthropological approach of living with the study population for 2.5 years, learning the language, and gaining subject confidence, 4) economic data derived from direct measurement rather than from interviews, and 5) the opportunity to measure numerous covariates seldom simultaneously measured in large surveys and therefore not adequately tested in previous studies. In our estimation, these advantages more than offset the benefits of standard demographic survey methods: large sample size and the possibility of drawing inferences to larger populations.

If one wishes to identify a woman from the study village who is expected to have a relatively short conception wait, then, according to these results, one should choose a young woman in a recent union with a young man. She should have had many pregnancies for her age and been demonstrably fecund in her recent past, as evidenced by the fact that she is nursing. One should pay less heed to the following: nulliparity vs. multiparity, duration of postpartum amenorrhea, sex of her last child, nutritional status, economic status, polygyny, and marital status (fiancée vs. married).

Perhaps the most striking result of this study is that, although we controlled for many different variables, heterogeneity in fecundability captured by the proxy variables gravidity and nursing status was substantial. The nature of this heterogeneity should be the subject of future investigations but might include genetic/ontogenetic influences not yet identified by human biologists. Familial patterns of fertility and subtle reproductive health risks over the life span are two new research directions. Personality differences might influence coital frequency and also play a role. For nearly two decades, debate has narrowly focused on a small number of parameters, such as nutrition and lactation, while other variables remain to be discovered. To elucidate these variables, comprehensive high quality data sets will be needed.

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