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Socio-Economic Factors, Fertility and its Proximate Determinants in the Côte d'Ivoire

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A structural model of fertility and its proximate determinants is estimated for the Côte d'Ivoire. Female secondary schooling is found to raise the age of cohabitation and reduce the duration of breast-feeding. Age at cohabitation and the duration of breast-feeding are both shown to reduce fertility, implying that the two effects of female education identified are offsetting. Fro younger cohorts, the impact via age at cohabitation dominates. The assumption of exogeneity of cohabitation and breast-feeding to fertility is rejected and shown to under-state the extent to which these limit fertility.

1. Introduction[1]

This paper estimates a structural model of the socio-economic determinants of the fertility in the Côte d'Ivoire. Population growth in the country is high, even by the standards of sub-Saharan Africa; with an annual rate of growth of 3.8% from 1980–1991 (World Bank, 1993). Whether such population growth should be a cause for concern is a complex issue, with debate over whether it will reduce general living standards and, even if this were true, whether this justifies limiting population size (Boserup, 1985; Parfit, 1984). Somewhat side-stepping these questions, economists' normative assessments of population policy tend to focus on the divergence between private and social costs of fertility (for example, Birdsall, 1988). In the case of the Côte d'Ivoire, such divergences are likely to arise because of state subsidies for health, education and other services for children, together with the dangers of natural resource depletion[2]. These issues are not discussed further here, although they do provide a *prima facie* case for the importance of the determinants of fertility as an area of study.

Contraceptive prevalence is very low in the Côte d'Ivoire, so socio-economic variables are likely to cause variations in fertility mainly by affecting age at cohabitation and the duration of breast-feeding. This paper examines these relationships in two stages. First, the reduced form effects of socio-economic variables upon age at cohabitation and the duration of breast-feeding are modelled. Particular attention is focused upon the impact of female education, which tends to raise age at cohabitation and reduce the length of breast-feeding. The second stage estimates a 'reproduction function' capturing the effects of the two main 'proximate determinants' upon fertility. The possible endogeneity of the proximate determinants is controlled for by using the results of the first stage models to obtain consistent estimators.

The remainder of the paper is divided into four. Section II provides a brief discussion of the links between socio-economic factors, fertility and its proximate determinants. Econometric techniques and data are outlined in Section III. Empirical results are given in Section IV, beginning with reduced form estimates of the socio-economic factors influencing age at cohabitation and breast-feeding before proceeding to a structural model of the impact of these proximate determinants on fertility. Section V summarises and concludes.

2. Conceptual Framework

Micro-econometric studies of fertility typically adopt a reduced form approach, modelling fertility directly as a function of its socio-economic determinants (see Ainsworth, 1990, for such estimates for the Côte d'Ivoire)[3]. However, it is potentially more informative to estimate the structural relationships, explaining how socio-economic variables affect fertility. In particular, it possible to estimate the technological relationship (the 'reproduction function') between fertility and those variables ('inputs') which directly affect it[4]. The concept of inputs to a reproduction function is similar to that of the

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[1] This paper was written whilst the author was employed by the Centre for Study of African Economies, a Designated Research Centre of the Economic and Social Research Council. The data set used was made available by the World Bank to the CSAE, which the author acknowledges with thanks. The paper benefited from comments from Manuel Arellano, John Hoddinott and Matthew Lockwood.

[2] De-forestation is currently occurring at the rate of 6–16% in the Cote d'Ivoire; its forests may disappear in less than twenty years (World Bank, 1991). Population growth is likely to be one factor contributing towards this.

[3] Ainsworth used similar data to those analyzed here, namely, the Living Standards Survey for 1985. Female schooling was found to reduce significantly the number of children ever born whilst consumption per adult increased it.

[4] Reproduction functions were estimated by Rosenzweig and Schultz (1985b, 1987).

proximate determinants of fertility used by demographers[5]. These proximate determinants include the pattern of unions, the frequency of intercourse, pathological and natural sterility, breast-feeding, contraception, induced and spontaneous abortions. The data used here have information on only two of the proximate determinants of fertility: age at cohabitation and the duration of breast-feeding. Consequently, the following analysis may be subject to omitted variables bias. However, the pattern of marriages and the duration of breast-feeding are said to be the principal proximate determinants of fertility differentials in regimes of natural fertility in general (Bongaarts and Potter, 1983). Breast feeding may capture two effects: amenorrhea and abstinence after pregnancy. According to the 1980–81 Ivorian World Fertility Survey (WFS), mean durations of reported abstinence are longer than mean reported durations of amenorrhea (Republique de Côte d'Ivoire, 1984). Consequently, the duration of abstinence may often determine the non-susceptible period, not breast feeding. With cohabitation, marital instability is the main omitted factor. However, figures from the Ivorian WFS indicate that around 20% of women ever divorce and 80% remarry, so this factor may not be so important in the Côte d'Ivoire[6]. In a developed country, contraceptive use would be crucial but in Côte d'Ivoire, it is relatively unimportant. According to the WFS, only 3.8% of women used any method of family planning and only 0.6% used an effective method such as the pill or IUD. Although use of modern methods of contraception may have increased since then, World Bank Sector Reports indicate current rates of usage of only around 3% (Ainsworth, 1991)[7].

One problem in estimating a reproduction function is that many of the inputs to it may be endogenous. In particular, there may be unobservables which affect both fertility and its proximate determinants. This might arise if fertility were consciously controlled through cohabitation and breast-feeding decisions. In such circumstances, one source of endogeneity bias could be unobserved fecundity. Women who are particularly fecund will have higher fertility, ceteris paribus, but may seek to offset this by choosing to breast-feed for longer. In such cases, treating breast-feeding as exogenous will tend to give estimates which under-state its effectiveness in limiting fertility. Simultaneity between fertility and its proximate determinants can arise even if couples do not consciously control their fertility. For example, Schultz (1986) notes that more fecund couples are more likely to have pre-marital conceptions and this may induce them to marry. This paper tests for the endogeneity of the duration of breast-feeding and the age at cohabitation using reduced form models of the socio-economic factors influencing them. The proximate determinants are modelled as functions of explanatory variables which are assumed both to be exogenous to fertility and not to directly affect fertility. Consequently, the predicted rather than actual values of the proximate determinants can be used to obtain consistent estimates of their effects upon fertility.

The reduced form approach to modelling the impact of socio-economic factors on the proximate determinants of fertility allows one to remain eclectic about the theory underlying the relationships involved. A woman's age at cohabitation is modelled as a function of her education and other personal characteristics, along with her parental background and some characteristics of her location such as local wage rates and social service infrastructure. A richer set of variables is used to explain the duration of breast-feeding, including current household characteristics, notably household consumption per capita. These variables were excluded from the models of age at cohabitation since they are simultaneously determined with it. Interpreting the reduced form results is problematic and there exist a number of competing conceptual frameworks. The dominant economic theory of fertility, sometimes termed the 'Chicago-Colombia' approach or the 'New Home (or Household) Economics',

[5] The literature on the proximate determinants of fertility originates from Davis and Blake (1956). They introduced the notion of intermediate fertility variables through which, and only through which, socio-economic and cultural factors could affect fertility.

[6] I am grateful to Matthew Lockwood for making these points.

[7] Low usage of modern contraception may be due to lack of availability rather than individual choice. Ainsworth (1990) reports that there were virtually no family planning services available except from private sources in Abidjan.

focuses on demand for children (see Willis, 1973). In the absence of modern contraception, decisions about age at cohabitation and the duration of breast-feeding could be viewed as being derived from the demand for children. However, viewing fertility as the outcome of rational choice remains controversial, particularly in developing countries where social taboos and mores are strong (Lockwood, 1990)[8]. Moreover, even in a world of ultra-rationality, desired family size is unlikely to be the sole — or perhaps major — consideration in determining choices about cohabitation and breast-feeding. Cohabitation behaviour may also reflect choices about the timing of fertility, specialisation in different types of production (Becker, 1973) and the general attractiveness of offers received at different times (Boulier and Rosenzweig, 1984). Breast-feeding decisions may be influenced by the possible consequences for the health of infants. They will also reflect the trade-off between time and monetary costs. Breast-milk substitutes tend to be more expensive in pecuniary terms whereas breast-feeding may place constraints on the mother's time use[9].

3. Econometric Models and Data

The data used draws on a sample of women surveyed in the 1986 Côte d'Ivoire Living Standards Survey (CILSS). The CILSS was a large general purpose survey with a nationally representative sample. One woman aged sixteen or over was randomly selected from each household to answer the fertility section of the questionnaire. This subset of women provides the basic sample used throughout this paper. Information was gathered on the age of the women when they first began cohabiting, how long they breast-fed their most recent offspring and the number of children ever born. This section provides a brief description of the econometric techniques used to model these three variables.

3.1 Modelling Age at Cohabitation

Age at cohabitation was subject to right censoring: of the 1439 members of the Ivorian sample used in this paper, 213 had not cohabited by the time of survey. Figure 1 presents the data using the Kaplan-Meier hazard rate, which gives the number of women who began cohabiting at a particular age as a proportion of the number at risk of doing so. The empirical hazard does not follow a well-behaved pattern, with four peaks. The limited range of ages at which most women begin to cohabit leads to a large number of 'ties' in the data. There is heavy bunching of age at cohabitation around the years 15 to 20. Under 10% of women are estimated to have married or cohabited before the age of 15 but 80% will have done so before they are 21. Indeed, the data suggests that, on present trends, almost all Ivorian women will cohabit eventually. According to the empirical survivor function, only 5% of women would reach the age of 27 without having cohabited. This implies that age at cohabitation is likely to reflect timing issues: when to cohabit rather than whether to ever do so.

The right censoring of the data on age at cohabitation can be easily accommodated by focusing on the 'hazard' of cohabitation rather than age at cohabitation itself. If T_i is time until cohabitation, the hazard for individual i at time t is defined as:

$$h_i(t) = \lim_{\delta \to 0^+} \frac{Pr[t+\delta | T_i \geq t]}{\delta} \qquad (1)$$

[8] A related argument is that individuals operate according to 'bounded rationality', which in this context means maximising subject to social norms (McNicoll, 1980).

[9] Once again the limitations of the rational choice paradigm in this context should be noted. See Page and Lesthaeghe (1981) and Lesthaeghe (1989), chapters 1 and 3.

Explanatory variables are assumed to have proportional effects upon the hazard, so that:

(2) $$h_i(t) = h_0(t).\exp(\beta' X_i)$$

where X_i is a vector of explanatory variables, ß a vector of coefficients to be estimated and $h_0(t)$ is the baseline hazard. Two alternative approaches were adopted to model the baseline hazard. One is parametric, the Weibull model, whereby:

(3) $$h_0(t) = p.t^{p-1}$$

The second approach is to leave the baseline hazards unspecified. This is attractive because economic theory provides little guidance as to the appropriate parametric form for a baseline hazard. Furthermore, imposing an inappropriate distribution would cause the coefficients of the covariates to be inconsistently estimated. Moreover, the irregular and multiple peaked empirical hazard in Figure 1 contrasts with the smooth monotonic hazard of the Weibull model. This suggests that there is likely to be a large gain from not assuming a particular form for the baseline hazard in modelling the determinants of age at cohabitation.

Cox (1972) provides a partial likelihood method for estimating such semi-parametric proportional hazards model but this encounters severe difficulties in the presence of heavy ties in the duration data (Cox and Oates, 1984). Conversely, heavy ties increase the reliability of Meyer's (1990) method which treats the baseline hazards as nuisance parameters to be estimated. Since heavy ties are a feature of the data, Meyer's method is used here[10]. In particular, equation (2) can be manipulated to give:

(4) $$Pr[T_i \geq t+1 \mid T_i \geq t] = \exp[-\exp(\beta' X_i + \mu_t)]$$

where:

$$\mu_t = \ln[\int_t^{t+1} h_o(u)du]$$

Meyer's approach is to regard the μ_t as parameters and to estimate them directly using maximum likelihood techniques[11]. The probability of surviving until $T_i = k_i$, is the product of terms of the form of equation (4), for $t=1,k_i-1$. Hence, with N individuals, the log-likelihood function is:

(5) $$L(\mu,\beta) = \sum_{i=1}^{N} \{\tau_i \ln[1-\exp\{-\exp(\beta' X_i + \mu_{k_i})\}] - \sum_{t=1}^{k_i-1} \exp(\beta' X_i + \mu_t)\}$$

[10] Few cohabitations began when women were outside the ages of 12 and 26, so baseline hazards for cohabitation outside this age range were not estimated. Instead, those who cohabit beyond the age of 26 are treated as being right censored at 26 and the seven individuals who cohabited below the age of 11 were treated as if they cohabited at age 11.

[11] The mathematical optimisation routine used was NAG routine E04KDF, a modified Newton algorithm. The first derivatives were programmed by the author. Analytic expressions are given in Appleton (1992).

where C_i is the censoring time (the first period at which the individual would not be observed), $k_i = \min[T_i, C_i]$ and τ_i is a zero-one indicator for being uncensored, ie:

$$\tau_i \begin{array}{ll} = 1 & \text{if } T_i < C_i \\ = 0 & \text{else} \end{array}$$

Omitted heterogeneity can also be controlled for by modelling it as a multiplicative random variable, ε_i, so that the hazard becomes:

$$h_i(t) = \varepsilon_i h_0(t).\exp(\beta' X_i) \tag{6}$$

If ε_i follows a gamma distribution with mean one (a normalisation) and variance σ^2, then the log-likelihood becomes:

$$\begin{aligned} L(\mu,\beta,\sigma) = {} & \sum_{i=1}^{N} \ln\{[1+\sigma^2.\sum_{t=1}^{k_i-1} \exp(\beta' X_i + \mu_t)]^{-\sigma^{-2}} \\ & - \tau_i[1+\sigma^2.\sum_{t=1}^{k_i} \exp(\beta' X_i + \mu_t)]^{-\sigma^{-2}}\} \end{aligned} \tag{7}$$

3.2 Modelling the Duration of Breast-feeding

The duration of breast-feeding was analyzed using the sub-sample of women with at least one live birth. These women had been asked how long they had breast-fed their last child. Given the inclusion of old women in the sample, answers to this question are likely to be subject to considerable measurement error[12]. Conversely, the presence of mothers with recent births leads to pervasive right censoring covering 28% of all observations[13]. Once again, use of the tools of empirical survival analysis provide insights into the nature of underlying data. The empirical survivor function implies that only 2.4% of women stop breast-feeding before their child is six months old. This is encouraging, given the consensus of medical opinion that breast-feeding less than this length is unhealthy for the child (WHO, 1981). However, breast-feeding in Côte d'Ivoire is a lengthy business: almost 90% of women are predicted to continue beyond a year and on most measures of central tendency, 18 months is the typical stopping point[14]. There are three marked peaks in the data. Of the uncensored observations, 15% stop at a year; 23% at 18 months; and 23% at two years. Such peaks probably reflect women with hazy memories reporting the duration of breast-feeding in half-years.

This 'heaping' feature of the data was incorporated into the econometric models used to analyze the duration of breast-feeding. The starting point was the classical linear regression model:

$$y_i = \beta' X_i + u_i$$

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[12] The earliest instance of breast-feeding analyzed in the sample was of a child born in 1929. Lesthaeghe and Page (1980) note the likelihood of recall error on breast-feeding: ceasing breast-feeding is not a major event like a birthday or a death. A further problem is that respondents might retrospectively amend their reports so as to conform with norms of what is or is not proper (Trussell, van de Walle and van de Walle, 1990).

[13] The questionnaire did not identify whether respondents were still breast-feeding and consequently this had to be inferred. Where the age of the last child differed from the reported duration of breast-feeding by one month or less, it was assumed that the child was still being breast-fed.

[14] The health benefits of such extended breast-feeding are questionable (see Appleton, 1992).

where y_i is the duration of breast-feeding for individual i, X_i is a vector of explanatory variables with an associated vector of coefficients ß and u_i is an error term, iid _ $N(0,\sigma^2)$.

Right censoring and possible grouping of the data were allowed for by use of maximum likelihood techniques. Observations known to have been breast-fed for k_i months contribute the probability of this to the likelihood function, that is to say:

$$(8) \quad \begin{aligned} Pr(y_i = k_i) &= \phi\left(\frac{k_i - \beta' X_i}{\sigma}\right) \\ &= \frac{1}{\sqrt{(2\pi\sigma)}}\exp\left(-\frac{1}{2}\left(\frac{k_i - \beta' X_i}{\sigma}\right)^2\right) \end{aligned}$$

Observations right censored at month k_i contribute the following to the likelihood:

$$(9) \quad \begin{aligned} Pr(y_i > k_i) &= Pr(u_i > k_i - \beta' X_i) \\ &= 1 - \Phi\left(\frac{k_i - \beta' X_i}{\sigma}\right) \end{aligned}$$

Two versions of the model were estimated because of the bunching of responses at 12, 18 and 24 months. One takes the responses were genuine, contributing probabilities as given in equation (8). The other assumes that such observations reflect grouping of the data. In particular, the 'grouped' version of the model assumes that all that can be inferred from such observations is that breast-feeding ended within a five month interval around the reported value. Consequently, observations bunched at k_i contribute the following to the log-likelihood:

$$(10) \quad \begin{aligned} Pr(k_i+2 > y_i > k_i-3) &= Pr(ki+2-\beta' X_i > u_i > k_i-3-\beta' X_i) \\ &= \Phi\left(\frac{k_i+2-\beta' X_i}{\sigma}\right) - \Phi\left(\frac{k_i-3-\beta' X_i}{\sigma}\right) \end{aligned}$$

The log-likelihood function formed is thus:

$$(11) \quad \begin{aligned} L = {} & \frac{N_1}{2}\ln(2\pi^2) - \frac{N_1}{2}\ln(\sigma^2) + \sum\nolimits_{i=1,N} D_{i2}\ln\left(1 - \Phi\left(\frac{k_i-\beta' X_i}{\sigma}\right)\right) \\ & + \sum\nolimits_{i=1,N}\left(D_{i3}\ln\left(\Phi\left(\frac{k_i+2-\beta' X_i}{\sigma}\right) - \Phi\left(\frac{k_i-3-\beta' X_i}{\sigma}\right)\right) - \frac{D_{i1}(k_i-\beta' X_i)^2}{2\sigma^2}\right) \end{aligned}$$

where D_{i1} = 1 if uncensored, 0 else
D_{i2} = 1 if right censored, 0 else
D_{i3} = 1 if grouped, 0 else

and there are N observations, of which N_1 are uncensored.

3.3 *Modelling the Number of Children Ever Born*

The number of children ever born to a woman was modelled as a function of her cohabitation and breast-feeding behaviour. However, the reproduction function is complicated by the fact that 17% of women in the sample had never given birth. In such cases, it makes no sense to try to estimate the relationship the impact of breast-feeding on fertility. More generally, it is likely that the factors affecting the arrival of a first born may differ from those affecting subsequent births[15]. Consequently, fertility was modelled in two parts. First, a probit was estimated for the probability of a woman having given birth at least once. Then a regression model for the number of children ever born was estimated over the sub-sample of those women with at least one birth. Possible sample selection effects were controlled for following Heckman (1979). The potential endogeneity of age at cohabitation and the duration of breast-feeding with respect to fertility was explored using non-linear two stage least squares. Use of predicted values of these proximate determinants obtained from their reduced form models should provide consistent estimates if the assumption of exogeneity was false. However, such estimates would be inefficient under exogeneity, allowing tests of the assumption following Hausman (1978).

4. Results

4.1 Socio-economic influences on the proximate determinants of fertility

Tables 1 and 2 give definitions and descriptive statistics for the explanatory variables used to capture socio-economic influences on age at cohabitation and the duration of breast-feeding in the Côte d'Ivoire. Tables 3 and 4 report the estimates of the effects of these hypothesised determinants[16]. The results for the duration of breast-feeding were not sensitive to whether the heaped responses in the data were assumed to be genuine or not. However, the analysis of age at cohabitation did vary according to which of three hazard rate models was adopted.

Likelihood ratio tests at 5% significance rejected both the restrictions imposed upon the baseline hazard by the Weibull model and the assumption of no unobserved heterogeneity in the semi-parametric model[17]. The intuition behind these results can be seen from the comparison of the empirical hazard in Figure 1 with the predicted hazards from the three models evaluated at the means of all explanatory variables. The smooth pattern predicted by the Weibull model in Figure 3 is unable to match the spiked appearance of the empirical hazard, in contrast to the hazards of the semi-parametric models[18]. Figure 2 reveals the importance of allowing for unobserved heterogeneity, which leads to much stronger positive duration dependence. For example, the baseline probability of a 26 year old who has never

[15] Sprague (1990) suggests that the decision to have a first child is quite different from the decisions to have subsequent children because the latter involve birth spacing issues.

[16] The explanatory variables included in the final forms were selected using a general-to-specific approach, with sequential exclusion of the least significant explanatory variables until all those remaining had a t-ratio of more than one.

[17] The LR test statistic for the Weibull parameterisation is 256.9, far greater than 26.1, the critical of the chi-squared distribution at 5% significance with 14 degrees of freedom. The LR test statistic for no unobserved heterogeneity is 18.6, in excess of the 3.84, the critical value of the chi-squared with one degree of freedom.

[18] A logistic parametric model is also plotted, since unlike the Weibull model, this does not constrain the baseline hazard to be monotonic. However, most attention is focused on the Weibull model since, being a proportional hazards model, it is most comparable with the semi-parametric estimates.

cohabited starting to do so is nearly 100% assuming mean heterogeneity. Without allowing for omitted heterogeneity, the probability is only 60%[19].

It is interesting to compare the gains in allowing for a flexible form for the baseline hazard with those from subsequently controlling for omitted heterogeneity. The former generates far larger improvements in overall explanatory power as measured by the value of the log-likelihood but is less important in producing unbiased estimates of the effects of covariates. Moving from the Weibull to the semi-parametric specification does not lead to any sign reversals, nor does it change which variables are significant at 10% in both models. Moreover, the size of the estimated coefficients are also roughly similar, although there are some non-negligible differences. By contrast, controlling for unobserved heterogeneity leads to large and unsystematic changes in the size and significance of the coefficients on observed determinants of age at cohabitation[20]. We now discuss the estimated effect of the explanatory variables, focusing upon the results of the semi-parametric model with heterogeneity since this is the most general model.

Female secondary schooling had powerful effects in raising the age at cohabitation and decreasing the duration of breast-feeding. One grade of schooling reduced the hazard of cohabitation by 36% in the case of upper secondary schooling, 13% in the case of lower secondary schooling but only 4% in the case of primary schooling (the latter being insignificant at the 10% level)[21]. A woman with no schooling but mean values for other characteristics has an expected age at cohabitation of 17.4; her counterpart with full secondary schooling has an expected age at cohabitation of 21.5. At the means of all regressors, a woman with full secondary schooling is likely to breast-feed for just under a year compared with one and a half years for an uneducated mother, and around one and a third years for a primary graduate[22]. Variables for the mother's endowments other than education were generally insignificant. These included maternal height together with various proxies for her parental background, both occupation and education. However, the insignificance of dummy variables for parental background may be due to low sample variation: the overwhelming majority of the women in the sample came from uneducated peasant households.

A common interpretation of the effect of female schooling is that it operates by raising the opportunity cost of women's time. However, no effects were found for a direct proxy for the opportunity cost of female time in rural areas, cluster specific wage rates for women's agricultural labour[23]. One should be cautious about inferring too much from this, since there

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[19] The latter figure is much higher than the empirical hazard plotted in Figure 1 because marriage observed at later ages do not occur for women with mean unobserved heterogeneity. In particular, those who are less likely to cohabit for unobserved reasons are more likely to survive to such late ages without having cohabited. A similar argument in the context of unemployment spells was emphasised by Lancaster (1979).

[20] A striking example of the sensitivity of the estimated coefficients to allowance for heterogeneity is the estimated effect of female secondary schooling. Assuming no heterogeneity, there is no difference in the effects of grades of upper and lower secondary schooling; with allowance for heterogeneity, the coefficient on the former increases by a factor of three.

[21] Part of this reflects an incompatibility between schooling and marriage; a dummy variable for whether the woman was still at school has a negative and highly significant effect on age at cohabitation.

[22] An inverse link between female education and breast-feeding is commonly found: see Wolfe and Behrman (1982), Akin et al (1985), Oni (1985) and Barrera (1991).

[23] Barrera (1991) found the community wage for female labour to have a significant negative effect in the Philippines.

is very little female labour market participation in rural Côte d'Ivoire (Appleton, Collier and Horsnell, 1990). Certainly, age at cohabitation is predicted to be higher in urban areas, where female employment is more common. However, breast-feeding is not predicted to be significantly shorter in urban areas. Moreover, where women have a higher share of household cash income, they tend to breast-feed for longer[24]. This might seem perverse if women's share of household cash income reflected the opportunity cost of women's time. Instead, it may be that, where women have more economic leverage within the household, they are able to increase the spacing of their births through abstinence and longer breast-feeding[25]. Perhaps the only good evidence of an effect on breast-feeding of the value of women's time is distance of the household to its main drinking water source. Since collecting water is women's work in Côte d'Ivoire, distant water sources will increase the value of their time, ceteris paribus. The models predict that an increase in distance to drinking water of one standard deviation will reduce the duration of breast-feeding by three weeks. This suggests that a higher opportunity cost of women's time increases their fertility, contrary to the negative relation conventionally hypothesised.

Female education may influence the proximate determinants of fertility via an income effect. Income was not controlled for in the model of age at cohabitation, since it could be regarded as simultaneously determined. However, household consumption per capita was entered in the models of breast-feeding and had a strong negative effect[26]. This may reflect a derived demand for children or the ability to afford costly breast-milk substitutes. However, the price of rice, was wholly insignificant and rejected from the final form of the equation, despite a rice-water gruel being a common weaning food in the Côte d'Ivoire (Schwartz, 1978). This may partly reflect historical and other sources of measurement error, although similar results have been found in studies of the Philippines (see Akin et al (1985); Barrera (1991)[27].

Factors affecting child 'quality' and child quantity may affect desired fertility and hence its proximate determinants (Becker and Lewis, 1973). However, there is little evidence of consistent effects of such factors on age at cohabitation and the duration of breast-feeding. Where the returns to child labour are high, one might expect higher desired family size and thus earlier cohabitation and shorter breast-feeding. However, only one of these predictions - younger cohabitation — receives support. In particular, in clusters where children under 12 worked, the model predicts a 23% increase in the hazard of cohabitation *ceteris paribus*. Conversely, breast-feeding is predicted to last significantly longer in such clusters and also to rise with the cluster wage for child agricultural labour. Similarly, indicators of local school availability and quality — indicators of the price of child quality — have mixed effects on the two proximate determinants. No evidence was found that women breast-feed boys for longer than girls, with a dummy variable for a female child being positive but wholly

[24] The predicted value of this variable was used after a Hausman test rejected its exogeneity (Hausman, 1978). The author is grateful to John Hoddinott for the provision of this variable. Details of its construction and the instruments used are given in Hoddinott and Haddad (forthcoming).

[25] The author is grateful to Matthew Lockwood for drawing his attention to this point.

[26] Consumption per capita was used as a measure of household wealth. A Hausman test, identified by household durables and other assets, did not reject the assumption that consumption per capita was exogenous.

[27] Akin et al (1985) also found the price and availability of formula foods to have no effect on the duration of breast-feeding, although distance to the nearest supplier did have a significant positive effect.

insignificant in preliminary estimates[28].

Health-related variables affect the duration of breast-feeding but not age at cohabitation. In particular, women who use non-natural sources of drinking water are likely to wean their children earlier. This may be because they believe such water to be safer[29]. Dummy variables for all three man-made water sources are significant, with coefficients whose sizes accord with what one might expect to be the perceived purity of the source: namely, piped water reduces breast-feeding the most, followed by water purchased from a vendor or obtained from a water truck, and then by water from a well. Living close to local medical facilities appears to increase the length of breast-feeding. This may reflect encouragement of breast-feeding by health practitioners, who are likely to have more influence on people nearby.

4.2 Estimates of a two stage reproduction function

Table 5 gives definitions and descriptive statistics for the inputs to the reproduction function. Table 6 reports estimates of the two-stage reproduction function. The first stage gives probit estimates of the probability of having a child, with the woman's age and cohabitation behaviour used as explanatory variables. Cohabitation was represented by a one-zero indicator for having ever cohabited and by a quadratic for age at cohabitation (set equal to current age where the woman had never cohabited). The second stage uses ordinary least squares regressions for the determinants of children ever born across a sample of women who had given birth to at least one child[30]. At each stage, three alternative estimates are presented to investigate the possible endogeneity of cohabitation and — in the second stage — breast-feeding with respect to fertility. When exogeneity is assumed, actual behaviour is used to explain fertility; when endogeneity is allowed for, use of predicted behaviour as regressors will give consistent estimates[31]. Use of both actual and predicted behaviour permits Hausman tests of exogeneity. At both stages of the reproduction function, Hausman tests rejects the assumption of exogeneity at 10% and hence attention is focused upon the results using predicted behaviour[32].

Age at cohabitation strongly reduces both the probability of giving birth and the number of

[28] This contrasts with Barrera's (1991) positive coefficient for male gender in the Philippines, which was almost significant at 10%.

[29] However, estimates of the determinants of morbidity on the same data cast doubt on whether such beliefs would be justified (Appleton, 1992).

[30] A Heckman two stage sample selection correction term used in preliminary estimates was wholly insignificant. This may be because it was identified only by functional form.

[31] The results of Tables 3 were used to calculate a predicted probability for an individual cohabiting and a predicted age at cohabitation. If predicted age at cohabitation exceeded an individual's age, the regressor 'predicted age at cohabitation' was set equal to current age. Individuals predicted to cohabit after age 27 were assumed to cohabit at age 31 (the mean for such observations who were uncensored).

[32] In the first stage, the relevant test uses the likelihood ratio and has a value of 7.7, compared with the critical value of the chi-squared distribution with three degrees of freedom of 7.82 at the 5% significance level and 6.25 at the 10% level. Given that the Hausman test is a low power test, these results suggest that age at cohabitation is endogenous to the probability of a first birth. In the second stage, the exogeneity of age at cohabitation and breast-feeding can be separately tested by referring to their t-ratios in the encompassing model. The exogeneity of both variables can be rejected independently at 1% significance.

children conditional on ever having given birth. The effects are very large[33]. Figure 4 shows the impact of age at cohabitation upon the probability of a twenty year-old giving birth. Values for age at cohabitation beyond twenty give the probability of a twenty year old having given birth if she had never cohabited. Consequently, the large drop in the plotted probability between the years of twenty and twenty-one provides one measure of the effect of cohabitation. The graph shows that, were a woman to start to cohabit at age twenty, this would increase the chances of her giving birth by 80 percentage points[34]. Age at cohabitation also raises the expected number of births conditional upon at least one having occurred. Delaying cohabitation by three years is predicted to reduce the number of children she has ever given birth to by one. Interestingly, increases in age at cohabitation has the same impact as an equivalent reduction in a woman's child-bearing years. There is no 'catch-up' effect of higher fertility at older ages among women who cohabited relatively late.

In both stages of the reproduction function, the effects of cohabitation are around twice as large in absolute sign as those estimated without controlling for endogeneity. When cohabitation is assumed to be exogenous to fertility, having cohabited increases the probability of giving birth by the age of twenty by only 48% rather than 80%. Similarly, the coefficient upon actual age at cohabitation over-estimates by a factor of two the postponement in age at cohabitation required to reduce the number of children ever born by one. Given the results of the Hausman test, this implies that standard estimates of the effects of cohabitation on fertility are subject to an positive endogeneity bias. Women who, for unobserved reasons, are more likely to cohabit late are also *ceteris paribus* more likely to give birth. Not allowing for this positive unobserved correlation underestimates of the effect of cohabitation upon fertility.

Breast-feeding only significantly limits fertility among women aged 35 to 55. This is surprising since the inhibiting biological effect of breast-feeding on fertility should hold for young as well as old. Part of the difference is presumably due to the fact that such Table 6 captures the cumulative impact of such effects, which will increase with the number of children born and hence with age. It may also be that there is an omitted variables problem. Younger women who tend to breast-feed for shorter periods — for example those with education — may use abstinence to reduce the risk of conception in the period immediately after weaning. Whatever the explanation, the effect of breast-feeding upon the fertility of older women is substantial. A one year increase in the duration of breast-feeding will reduce the number of children ever born to a woman of fifty-five by over four. Using actual rather than predicted breast-feeding generates an effect only half as powerful. As with the endogeneity biases on age at cohabitation, this indicates that women who tend to have more children behave in ways that are more likely to reduce their fertility[35].

Interpreting the existence of such endogeneity biases is not straightforward. The results contradict Schultz's (1986) hypothesis, whereby fecund women are assumed to be more likely to marry early. Instead, it may be that Ivorian women who have delayed cohabitation — primarily by remaining in school — tend to be healthier and more fertile. The biases on breast-feeding resemble findings by Rosenzweig and Schultz (1985a, 1985b) and Kelley and Schmidt (1985) that the effectiveness of contraceptive use in limiting fertility is underestimated when endogeneity is not controlled for. Fecund women may attempt to limit their fertility through prolonged breast-feeding. However, women are only likely to learn gradually about their fecundity, which may help to explain why breast-feeding has a

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[33] This is to be expected from the simple descriptive statistics. Only 3.2% of women who had given birth to a child had never cohabited. Conversely, only 6.8% of women who had cohabited had not given birth to a child.

[34] Delays in cohabitation prior to the age of twenty reduce the likelihood of having given birth. For example, Figure 4 shows that beginning to cohabit at twenty rather than fifteen reduces the risk of giving birth by the age of twenty by eighteen percentage points. Delaying cohabitation has a diminishing effect on the likelihood of giving birth. Postponements in the age of cohabitation beyond twenty do not significantly alter the probability of giving birth.

[35] This is result is particularly striking given that women who tend to cohabit later are very different from women who tend to breast-feed for long in observable ways — such as education.

powerful effect on fertility for older women only.

5. Summary and Conclusions

This paper illustrates a number of methodological problems in modelling fertility and its proximate determinants. The choice of the hazard rate model was found to affect estimates of the determinants of age at cohabitation, with control for unobserved heterogeneity being particularly important. By contrast, appropriate econometric techniques to allow for the 'heaped' feature of the data on duration of breast-feeding brought no new insights. Age at cohabitation and the duration of breast-feeding were both found to be endogenous to fertility. This throws doubt on demographic models of the proximate effects of fertility which assume exogeneity. The directions of the biases imply that women who have a tendency to have more children also marry later and breast-feed for longer. Assuming exogeneity would understate the effectiveness of increases in these proximate determinants in lowering fertility.

In the Côte d'Ivoire, female education had particularly powerful effects on the proximate determinants of fertility, increasing age at cohabitation but decreasing the duration of breast-feeding. Evaluating at the means of other variables, a woman with complete secondary schooling will cohabit four years later than one with no schooling and breast-feed for half a year less. However, both age at cohabitation and the duration of breast-feeding negatively affect fertility, although the effect of the latter was confined to older age groups. Hence, the effects of female education upon age at cohabitation and the duration of breast-feeding have opposite implications for fertility. Beginning to cohabit at 21 years of age rather than 17 would reduce the expected number of children (conditional upon at least one) of a 55 year old woman by 1.4. The corresponding increase from breast-feeding for 12 months rather than 18 would be 2.2. Consequently for older women, education may have had a positive net effect on fertility. Caution is required because almost no women that old had education. In the sample as a whole, female education will tend to be associated with lower fertility; the positive effect on age at cohabitation dominating the insignificant effects of the duration of breast-feeding on the fertility of younger cohorts.

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Table 1:
Variables used in Analysis of Age at Cohabitation

Variable	**Description**	**Mean**	**S.D.**
mage	years of age from 25 to 50	10.197	10.016
age50	years of age from 50 to 55	0.749	1.707
oage	years of age over 55	1.241	4.600
grp	grades of primary school	1.444	2.452
grls	grades of lower secondary school	0.356	1.046
grus	grades of upper secondary school	0.067	0.416
pupil	1 if student, 0 else	0.039	0.193
noclass	number of classes in local primary school *	3.633	3.509
secdist	distance to nearest secondary school (km) *	15.789	20.432
pmidist	distance to nearest PMI clinc (km)	22.346	34.761
kidswork	1 if children under 12 work in cluster, 0 else *	0.222	0.415
manwage	cluster wage for adult male agricultural labour (CFAF/day) *	345.830	349.775
mktdist	distance to nearest market (km)	2.102	4.190
primf	1 if father had some schooling, 0 else	0.081	0.272
tradef	1 if father a trader, 0 else	0.032	0.176
muslim	1 if Muslim, 0 else	0.331	0.470
othchris	1 if non-Catholic Christian, 0 else	0.095	0.293
nonivory	1 if non-Ivorian, 0 else	0.152	0.359
kru	1 if Kru, 0 else	0.120	0.324
voltaic	1 if Voltaic, 0 else	0.108	0.310
urban	1 if urban resident, 0 else	0.420	0.494
abidjan	1 if Abidjan resident, 0 else	0.198	0.399
ssav	1 if rural South Savannah residence, 0 else	0.207	0.405
nsav	1 if rural North Savannah reside residence, 0 else	0.071	0.257

Sample Size 1439.

* = unknown (set to zero) for urban clusters
S.D. = standard deviation

Table 2:
Variables used in Analysis of Duration of Breast-feeding

Variable	Description	Mean	S.D.
birthyr	year of birth of child (minus 1900)	78.091	10.526
grs	grades of secondary school	0.269	1.059
grp	grades of primary school	1.075	2.187
wdist	distance to drinking water (m)	188.919	342.602
aconpc	household consumption per capita excl. own offspring (1000 CFAF pa)	14.220	16.026
fridge	1 if household has a fridge, 0 else	0.208	0.406
latrine	1 if household has a pit latrine, 0 else	0.428	0.495
flush	1 if household has a flush toilet, 0 else	0.128	0.334
tap	1 if household drinks tap water, 0 else	0.180	0.384
buyw	1 if household buys water, 0 else	0.114	0.318
well	1 if household drinks well water, 0 else	0.570	0.495
noclass	number of classes in local primary school *	3.975	3.450
hdist	distance to nearest health facility (km) *	6.853	9.503
xmhdist	extra distance to nearest maternity hospital (km) *	2.893	9.336
kidwage	cluster wage for child agricultural labour (CFAF/day) *	284.890	294.688
urban	1 if urban resident, 0 else	0.363	0.481
abidjan	1 if Abidjan resident, 0 else	0.142	0.349
muslim	1 if Muslim, 0 else	0.337	0.473
epfinca	predicted female share of cash income	0.210	0.173

Sample Size 1052.

* = unknown (set to zero) for urban clusters
S.D. = standard deviation

Table 3:
Hazard Rate Models of Age at Cohabitation

	Semi-parametric With Heterogeneity			**Proportional Hazards Without Heterogeneity**			**Weibull Model**		
Variable	Coefficient			Coefficient			Coefficient		
mage	-0.011	(-1.66)	*	-.481E-02	(-1.16)		-.596E-02	(-1.46)	
age50	-0.143	(-3.09)	***	-0.069	(-2.62)	***	-0.069	(-2.48)	**
oage	-0.023	(-1.79)	*	-0.011	(-1.31)		-0.013	(-1.29)	
grp	-0.044	(-1.45)		-0.022	(-1.12)		-0.013	(-0.74)	
grs1	-0.130	(-1.69)	*	-0.098	(-1.92)	*	-0.112	(-2.54)	**
grs2	-0.361	(-2.10)	**	-0.101	(-1.05)		-0.154	(-1.47)	
pupil	-0.806	(-2.17)	**	-0.666	(-2.25)	**	-0.505	(-1.68)	*
noclass	-0.064	(-1.86)	*	-0.052	(-2.47)	**	-0.052	(-2.59)	***
secdist	-.805E-02	(-2.13)	**	-.388E-02	(-1.73)	*	-.375E-02	(-1.67)	*
pmidist	0.248E-02	(1.36)		0.144E-02	(1.27)		0.159E-02	(1.25)	
kidswork	0.234	(1.73)	*	0.147	(1.75)	*	0.142	(1.78)	*
manwage	0.287E-03	(1.07)		0.249E-03	(1.44)		0.278E-03	(1.49)	
mktdist	-0.036	(-2.38)	**	-0.024	(-2.73)	***	-0.026	(-3.32)	***
primf	-0.355	(-1.56)		-0.173	(-1.13)		-0.130	(-0.87)	
tradef	-0.320	(-1.17)		-0.292	(-1.73)	*	-0.360	(-2.59)	***
muslim	0.195	(1.66)	*	0.148	(2.03)	**	0.163	(2.41)	**
othchris	-0.183	(-1.08)		-0.092	(-0.86)		-0.052	(-0.53)	
nonivory	0.267	(1.86)	*	0.234	(2.61)	***	0.240	(2.78)	***
kru	0.461	(2.69)	***	0.207	(2.10)	**	0.189	(2.11)	**
voltaic	0.374	(1.70)	*	0.178	(1.34)		0.149	(1.27)	
urban	-0.809	(-2.48)	**	-0.581	(-2.95)	***	-0.572	(-3.08)	***
abidjan	-0.543	(-3.42)	***	-0.332	(-3.38)	***	-0.356	(-4.32)	***
ssav	-0.519	(-3.42)	***	-0.373	(-4.14)	***	-0.323	(-3.80)	***
nsav	-0.784	(-2.60)	***	-0.463	(-2.54)	**	-0.425	(-2.62)	***
sigma	1.070	(3.29)	**						
rho							2.664	(43.79)	**
Log-Likelihood	-3154.60			-3163.90			-3292.29		

t-ratios in brackets
sigma = heterogeneity variance
rho = Weibull scale parameter

Table 4:
Censored Regression Models of the Duration of Breast-feeding

	Assuming Heaped Responses			Assuming Genuine Responses		
Variable	**Coefficient**			**Coefficient**		
constant	27.726	(11.87)	**	27.358	(11.98)	**
birthyr	-0.068	(-3.22)	***	-0.065	(-3.18)	***
grs	-0.989	(-3.75)	***	-0.963	(-3.70)	***
grp	-0.190	(-1.38)		-0.184	(-1.36)	
wdist	-.232E-02	(-2.47)	**	-.226E-02	(-2.46)	**
aconpc	-0.033	(-2.21)	**	-0.033	(-2.25)	**
fridge	0.862	(1.56)		0.859	(1.59)	
latrine	-0.619	(-1.09)		-0.625	(-1.13)	
flush	1.851	(1.69)	*	1.774	(1.65)	*
tap	-4.380	(-3.65)	***	-4.289	(-3.66)	***
buyw	-3.690	(-2.88)	***	-3.608	(-2.87)	***
well	-1.541	(-1.70)	*	-1.532	(-1.72)	*
noclass	-0.226	(-1.51)		-0.235	(-1.61)	
hdist	-0.063	(-1.97)	**	-0.063	(-2.02)	**
xmhdist	-0.052	(-1.87)	*	-0.052	(-1.90)	*
kidwage	0.217E-02	(1.82)	*	0.204E-02	(1.75)	*
urban	-1.952	(-1.42)		-2.035	(-1.51)	
abidjan	-1.231	(-1.34)		-1.180	(-1.31)	
muslim	1.019	(2.08)	**	1.018	(2.13)	**
epfinca	2.644	(1.97)	**	2.523	(1.92)	*
variance	36.665	(19.50)	**	36.342	(19.85)	**
Log-Likelihood	-1767.87			-2508.27		

t-ratios in brackets

Table 5:
Variables Used in Reproduction Function Estimates

a) Definitions

CHILD	1 if given birth to live child, 0 else
NCHILD	number of children ever born
AGE35	age if aged 35 or under, 35 else
AGE55	age minus 35 if aged between 35 and 55; 0 if aged under 35; 20 if aged over 55
OAGE	age minus 55 if aged over 55, 0 else
COHAB	1 if ever cohabited, 0 else
DALONE20	age at cohabitation if cohabited at before age 21; 20 if cohabited after age 20; current age if never have cohabited
PCOHAB	probability ever cohabited on or before current age
PDALONE20	predicted value of DALONE20
PBFAGE55	predicted months breast-feed last child, multiplied by AGE55
FAGE55[36]	actual months breast-fed last child, multiplied by AGE55
AGE	age (years)
DBF	months breast-fed last child
PDBF	predicted months breast-fed last child
DALONE	age at cohabitation or current age if never have cohabited
PDALONE	predicted years until cohabitation (inclusive)

A '2' following a variable name indicates that it is a squared term.

b) Descriptive Statistics

	Sample 1		**Sample 2**	
	Mean	**Std.Dev.**	**Mean**	**Std.Dev.**
CHILD	0.832	0.374	1.000	0.000
NCHILD	4.190	3.386	5.034	3.086
AGE35	28.811	6.950	30.400	5.831
AGE55	5.330	7.507	6.030	7.691
OAGE	1.199	4.475	1.273	4.500
PCOHAB	0.874	0.199	0.930	0.104
PDALONE20	17.890	1.211	18.001	1.107
PDALONE202	321.500	43.451	325.260	40.267
COHAB	0.865	0.342	0.968	0.176
DALONE20	17.311	2.218	17.283	2.255
DALONE202	304.580	74.441	303.790	75.286
PBFAGE55	96.138	151.470	118.850	155.020
BFAGE55	93.781	156.440	116.020	161.230
AGE	35.339	15.517	37.703	14.706
DBF	13.111	11.201	17.564	5.688
PDBF	13.719	10.503	18.293	2.751
DALONE	17.963	3.582	17.938	3.654
PDALONE	17.984	1.445	18.093	1.346
Sample Size	1235		1028	

Sample 2 subsets those who in Sample 1 who have given birth
Std.Dev. = Standard Deviation

[36] If currently breast-feeding last child, the predicted duration of breast-feeding was used.

Table 6:
Reproduction Function Estimates

a: Probit for Having a Child

Variable	Model A Coefficient			Model B Coefficient			Model C Coefficient		
CONSTANT	22.798	(1.96)	**	-12.051	(-3.60)	***	15.012	(1.16)	
AGE35	0.358	(2.50)	**	0.718	(7.07)	***	0.546	(3.32)	***
AGE352	-.549E-02	(-2.17)	**	-0.012	(-6.33)	***	-.903E-02	(-3.12)	***
OAGE	-0.015	(-1.27)		-0.019	(-1.70)	*	-0.020	(-1.61)	
PCOHAB	2.902	(4.63)	***				1.645	(2.28)	**
PDALONE20	-3.130	(-2.42)	**				-2.922	(-2.04)	**
PDALONE202	0.082	(2.28)	**				0.080	(2.01)	**
COHAB				1.338	(7.82)	***	1.240	(6.96)	***
DALONE20				0.403	(1.12)		0.473	(1.36)	
DALONE202				-0.016	(-1.54)		-0.018	(-1.77)	*
Log-likelihood:	-348.00			-292.61			-289.50		
Intercept only	-558.32								

b: O.L.S. Regression for Number of Children Ever Born Conditional on One Birth

Variable	Model A Coefficient			Model B Coefficient			Model C Coefficient		
CONSTANT	2.650	(2.20)	**	-1.284	(-2.32)	**	2.510	(2.12)	**
AGE35	0.273	(16.57)	***	0.286	(17.32)	***	0.284	(17.44)	***
AGE55	0.439	(5.01)	***	0.094	(3.15)	***	0.442	(5.14)	***
OAGE	-0.075	(-1.54)		-0.156	(-3.38)	***	-0.074	(-1.54)	
OAGE2	0.511E-02	(3.22)	***	0.614E-02	(3.92)	***	0.502E-02	(3.21)	***
PBFAGE55	-0.018	(-4.13)	***				-0.019	(-4.20)	***
PDALONE	-0.355	(-5.54)	***				-0.232	(-3.51)	***
BFAGE55				-.860E-03	(-0.64)		0.575E-03	(0.42)	
DALONE				-0.154	(-7.40)	***	-0.135	(-6.25)	***
R-squared	0.410			0.416			0.431		
Mean Std. Error	2.378			2.364			2.336		

t-ratios in brackets
Model A assumes endogeneity; Model B assumes exogeneity and Model C tests these assumptions.

Figure 1: Age at cohabitation; Kaplan–Meier Empirical Hazard

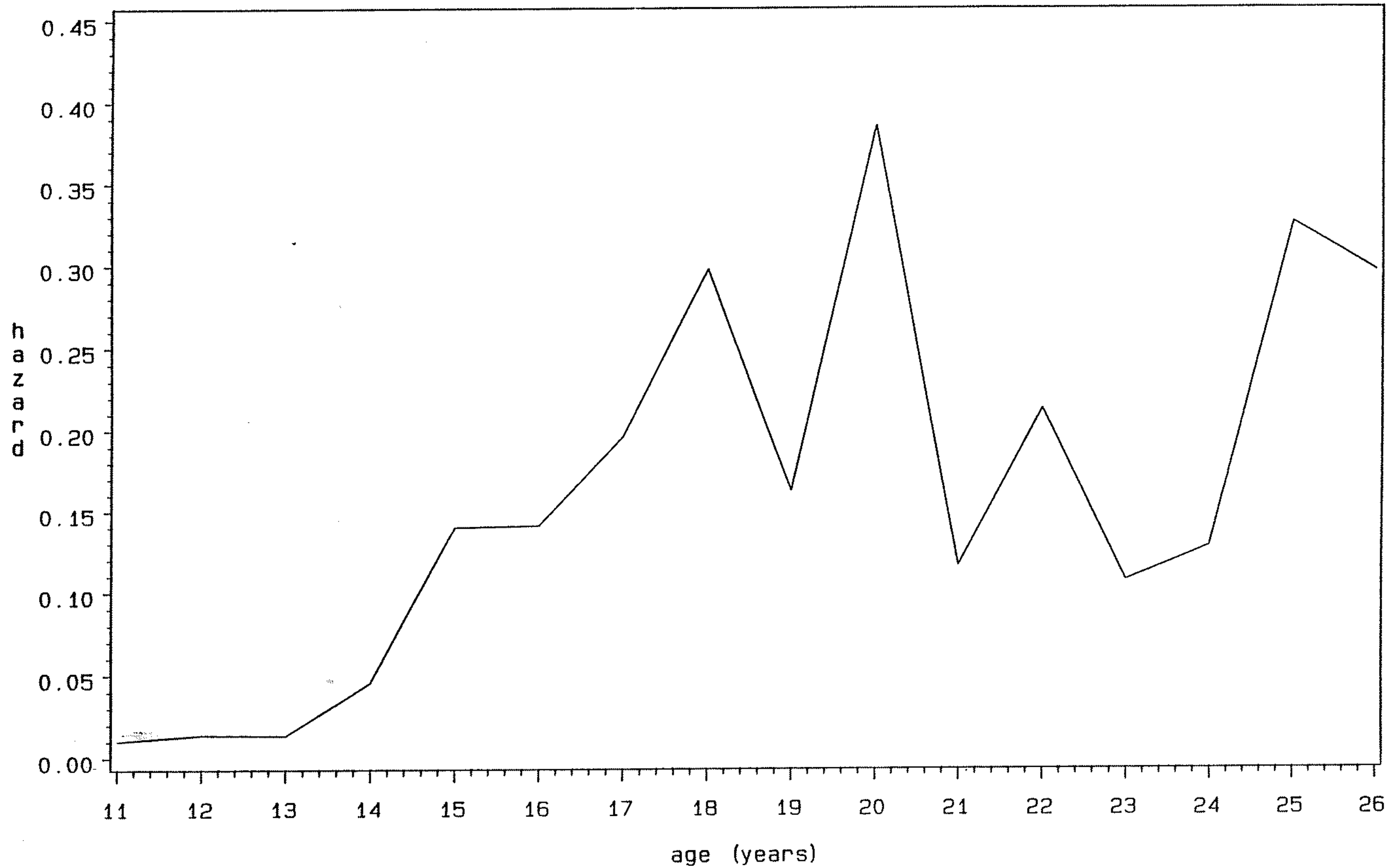
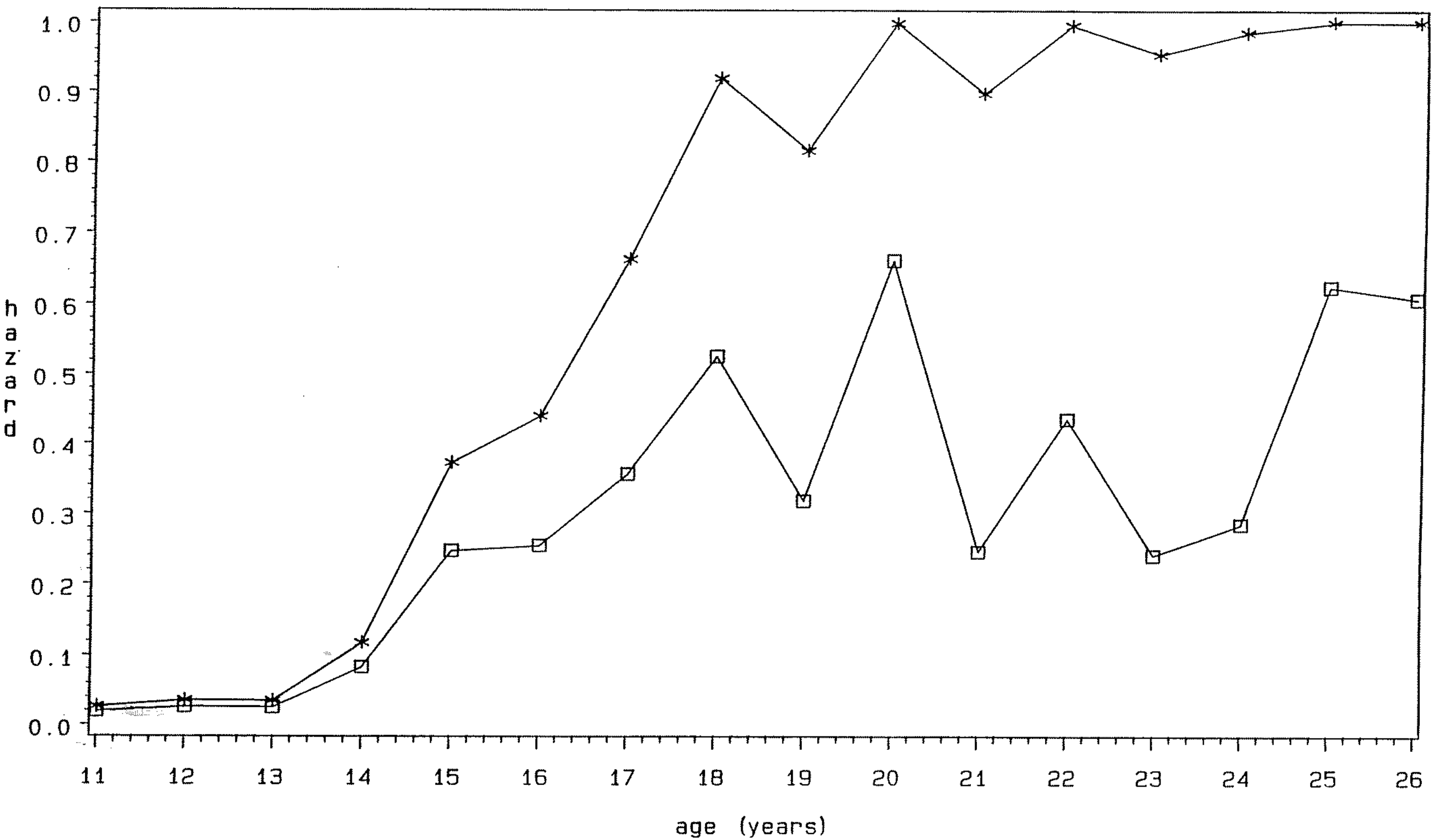


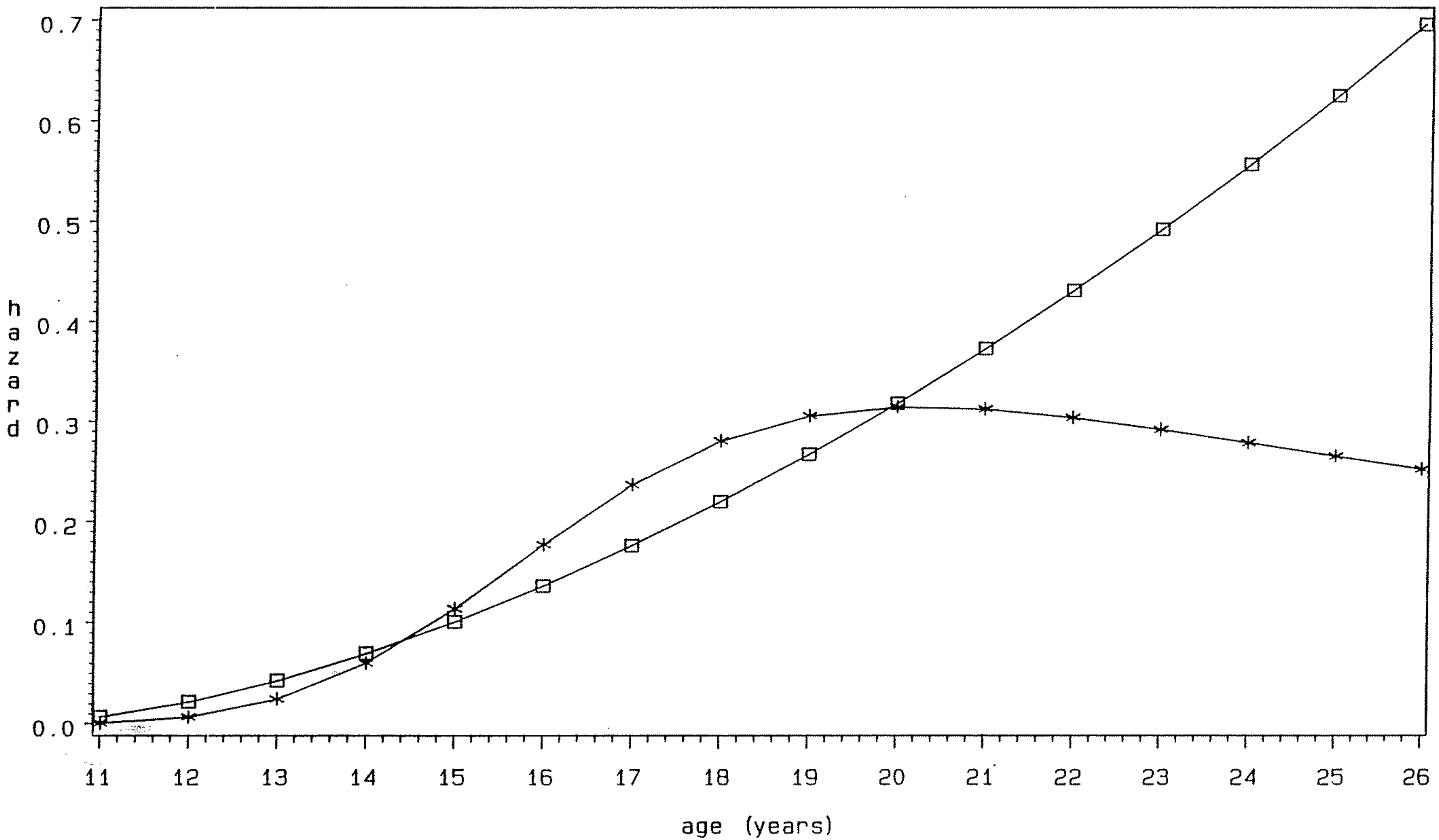
Figure 2: Age at Cohabitation; Semi-parametric Baseline Hazards

With and without Heterogeneity



square = no heterogeneity   star = with heterogeneity

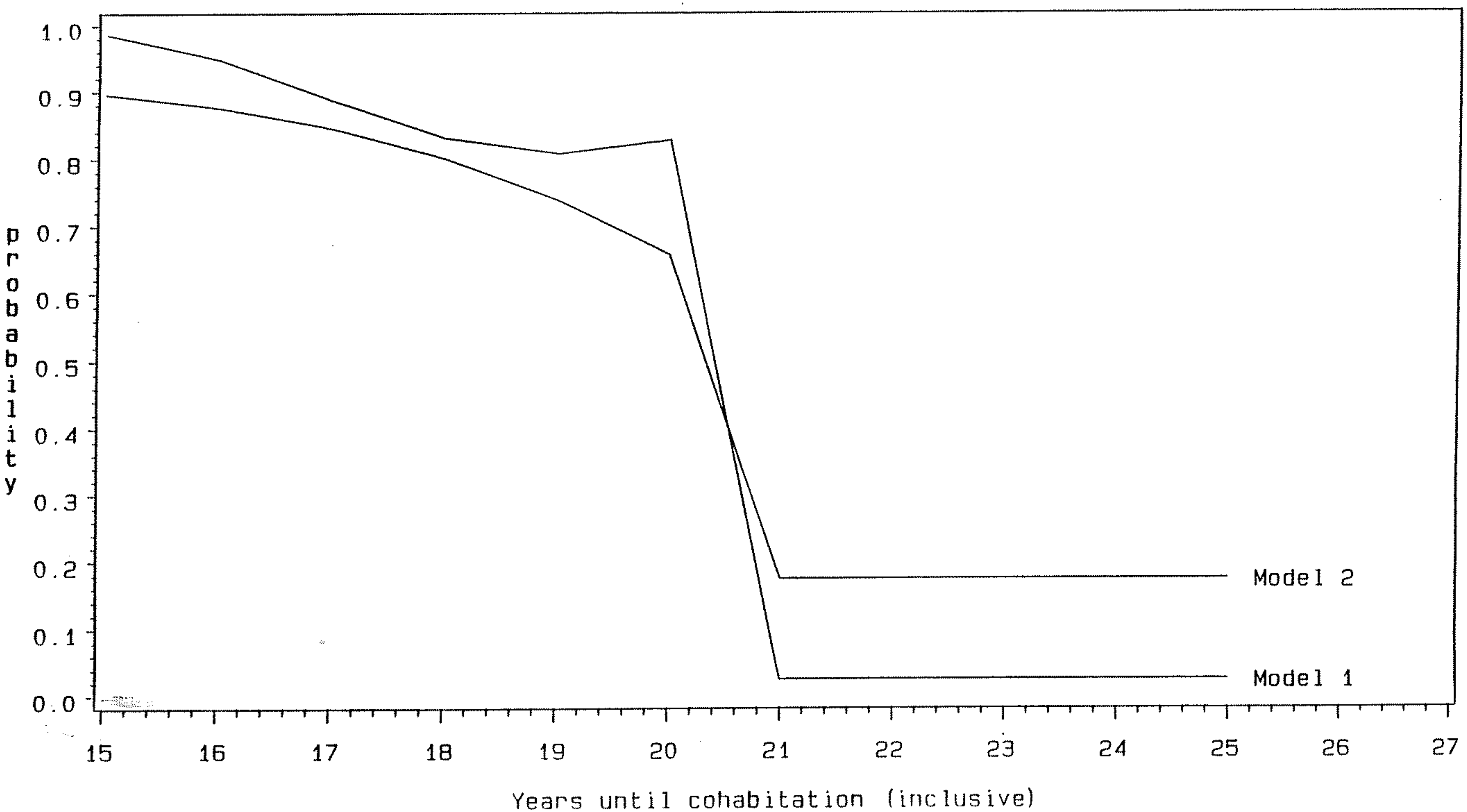
Figure 3: Age at Cohabitation; Parametric Hazards



square = weibull star = logistic
Evaluated at the mean of the reqressors

Figure 4: Age at Cohabitation and the Probability of Giving Birth

Probability given birth by age 20
Probit predictions: Women in Cote d'Ivoire 1986



Model 1 uses predicted age at cohabitation
Model 2 uses actual age at cohabitation