Journal of Quantitative Economics, Vol. 12, No. 1 (January 1996) 81 - 104

OLD-AGE SECURITY AND GENDER PREFERENCE HYPOTHESES : A DURATION ANALYSIS OF MALAYSIAN FAMILY LIFE SURVEY DATA*

LAKSHMI K. RAUT

University of Hawaii-Manoa, Department of Economics, Honolulu

This paper demonstrates that an accelerated hazard framework is more appropriate than commonly used proportional hazard framework to model the timing of marriage and timing and spacing of children. Using the 1976 Malaysian Family Life Survey Data and selecting an appropriate duration model, the paper tests the old-age security and gender preference hypotheses and estimates the replacement effect. The paper finds strong evidence for the old-age support hypothesis: If parents have sufficient wealth to support themselves during their old-age, they have longer birth intervals; the Chinese have higher likelihood of depending on their children for old-age support than the Malays, and higher is the husband's earnings, the less likely is the couple's dependence on children for such support. The paper finds weak evidence for son preference hypothesis: The number of sons has no effect on the birth intervals of the Malays whereas it has significant negative effect up to three children for the Chinese. Regarding the replacement effect the paper finds that in response to a child death, parents like to have shorter durations up to fourth child and thereafter, the effect is not significant; the desired family size of the Chinese is higher than that of the Malays. (JEL : C41)

1. INTRODUCTION

The old-age security and gender preference are two highly controversial hypotheses regarding the determinants of household fertility in less developed countries. The old-age security hypothesis postulates that in environments where parents face uncertainty about the ability to support themselves during old-age, they would expect such support from their children. This motive could be strong particularly in rural areas of less developed countries, where capital markets are imperfect and insurance markets are

* An earlier draft of the paper was presented at the Far Eastern Econometric Society Meeting, June 29-30, 1991, Seoul, Korea, Fifth Annual Meeting of the European Society for Population Economics, June 6-8, Pisa, Itlay and the Development Workshop at the University of California, Berkeley. Comments from Andrew Foster, Glenn Sueyoshi, and participants of the above meetings and workshop are gratefully acknowledged. The financial support for this research came from the grant RM 199-G, Committee on Research, University of California.

absent. Financial and physical assets that are available in these areas tend to yield low or negative interest rates, so that children may provide a more efficient hedge against old age disability risks. In Raut [1985, 1990, 1992]), I have examined the aggregate implications of old-age security hypothesis for population growth, capital accumulation, and income distribution of the subsequent generations. See Nerlove and Raut [1995] for a more recent account and Cain [1981, 1983], Nerlove, Razin and Sadka [1987], and Nugent [1985] for other references and further discussions on the old-age security hypothesis.

The gender preference hypothesis postulates that parents exhibit preference for having children of a particular gender. In most less developed societies, parents seem to have preference for sons rather than daughters. The preference for sons may be rooted purely in taste and cultural values or it could be the outcome of some economic calculations. For instance, sons generally stay with their parents while daughters are married off to another household, so that sons tend to provide better support in old-age as well as augment current household income. The extent to which the preference for sons occur as an economic response to underdeveloped capital markets and incomplete risk markets has been a long standing issue in economic demography literature. The consequences of this motive for population growth, and sex ratio have been examined by Ben-Porath and Welch [1976], Heer [1983], and Leung [1988] and the consequences on the allocation of human capital and bequest among children have been analyzed by Behrman, Pollak and Taubman [1982].

Empirical studies on the old-age security motive have been carried out using both cross country macro data and data from cross section of villages or counties, and the conclusions of these studies are controversial. For example, using cross country time series data to study the effect of introducing social security program on total fertility rate, Entwisle and Winegarden [1981], Hohm *et al* [1984] find evidence for old-age security hypothesis, whereas Kelly, Cutright and Hittle [1976] find no such evidence. Comparing the fertility rates of rural counties/villages with some type of old-age pension scheme with those without it, Cain [1981, 1983] (for India and Bangladesh), Nugent and Gillaspy [1983] (for Mexico) and Sanchez [1984] (for the U.S.) find evidence for the old-age security hypothesis, whereas Robinson [1986] finds no difference.[1] These studies do not use information on the developments of events over the life-cycles of couples which significantly affect fertility choices; we argue that the hazard rate approach is more appropriate in this context.

The empirical literature on the gender preference hypothesis has also documented controversial evidence (see Ben-Porath and Welch [1976] and Leung [1988] on this). The studies that relate the effect of number of sons to the subsequent birth intervals are the ordinary least square regression analyses of Ben-Porath and Welch [1976] using Bangladesh data, De Tray [1984] using Pakistan data and Leung [1988] using Chinese sample of the Family Life Survey Data. All these studies except Leung's are based on closed birth intervals and thus have sampling bias of throwing away the incomplete birth intervals. These studies do not control for important determinants of

1. See Nugent [1985] for an extensive survey of this literature.

birth intervals such as infant death, and child death. They find that the effect of number of sons on subsequent birth intervals is weakly positive for up to five children. In this paper I use an appropriate duration model and estimation procedure on a broader sample of households (as compared to Leung and others) and compare our findings with the previous literature. We improve our model and the estimation techniques along the following lines:

Fertility decisions are made sequentially by a couple over their life-cycles and thus are affected by the changes in socio-economic variables over their life-cycles. The perception about the degree of old-age insecurity, preference for son/daughter, and the occurrence of an infant or child death may depend at any time upon the number of surviving children, earnings profile of husband and wife, and stock of assets and therefore will vary over the life-cycles of a couple. Fertility decisions will also interact with the labor supply and savings decisions over the life-cycles of a couple. Empirical analyses that are based on completed fertility will not be able to capture these dynamic effects. Furthermore, the parameter estimates are subject to cohort and selectivity biases as a result of right censored data; for instance, the young women who have not completed their reproductive periods are dropped out of the right censored sample and thus the sample represents only those old women who survived until the survey date; their characteristics might have a systematic effect on fertility behaviour which is different from the behavior of a representative woman in the whole population. To circumvent some of these problems, we model fertility decisions as a sequential decision making process.

Empirical studies incorporating the sequential nature of fertility decisions are carried out by Heckman and Willis [1975], Newman and McCulloch [1984], Olsen and Wolpin [1982], Wolpin [1984], and Hotz and Miller [1988], Ben-Porath and Welch [1976]. Hotz and Miller studied the interaction between fertility and labor supply decisions over discrete-time life-cycle periods using longitudinal U.S. household survey data (more specifically, the Panel Survey of Income Dynamics data). Wolpin formulated a simplified model of life-cycle fertility decisions within a discrete time dynamic programming framework and was able to estimate the structural parameters of the model; Olsen and Wolpin used a waiting time regression framework to study the replacement effect. Both studies used the same dataset as ours. However, we use different model and estimation techniques and we shall contrast our findings with theirs. Ben-Porath and Welch used data from Bangladesh and applied an ordinary least square regression technique to estimate the effect of the number of sons on subsequent birth intervals.

Studies that come closer to the hazard rate approach of this paper are by Heckman and Willis and Newman and McCulloch. These papers, however, assume that a couple choose a contraception method to proportionately scale up or down the risks of a live-birth that would have resulted if the couple did not use any contraception. These type of models are known as proportional hazard models. The proportionality assumption imposes severe restrictions on fertility behavior and can lead to faulty inference about the effect of a regressor on the birth interval. For instance, consider two women: one is an unemployed housewife and does not wish to use contraception for timing her first birth. The probability of her giving a birth at time t given that she has not done so until t will be strictly positive for all t until she reaches her menopause. This probability is determined by biological law and is known as the base-line hazard rate. The second woman is identical to the first in all respects except that she is employed and likes to target her first birth with very small probability up to a point in time and thereafter likes to accelerate the probability of having the birth. It is clear that the life-cycle experiences of a couple directly affect the targeting of the age at marriage and birth intervals. The duration models that take into account these latter type of effects are known as accelerated hazard models and are more reasonable description of decisions regarding timing of marriage and timing and spacing of birth. I argue in the paper that when observed behavior follows an accelerated hazard model but we fit a proportional hazard model, we may end up with wrong inference about the effect of a covariate.

While there have been a few attempts to study the replacement and sex preference hypotheses in a duration framework, most of them have used proportional hazard models and no one, I believe, has used the duration framework to test the old-age security hypothesis directly. This framework has the advantages that it takes into account the sequential nature of fertility decisions, the stochastic nature of the reproductive process, right censoring of the sample and the effects of time varying measured and unmeasured heterogeneity. I use the accelerated hazard rate approach to study these hypotheses.

Section 2 models the timing of marriage and timing and spacing of births as an accelerated hazard model. Section 3 talks about econometric issues. Section 4 describes the data and defines the variables used in the paper. Section 5 reports the parameter estimates. Section 6 summarizes the results and concludes the paper.

2. PROPORTIONAL VS. ACCELERATED HAZARD MODEL OF TIMING OF MARRIAGE AND TIMING AND SPACING OF BIRTHS

There are several ways in which marriage and household fertility decisions are modeled in the literature. One approach, known as the household production framework (Becker [1965]), views that husband and wife use part of their time and goods purchased from the markets as inputs to produce non-marketed household goods and services such as children and child qualities. Parents derive utility from consumption of these household goods. The decisions regarding the timing of marriage and family size are jointly determined by balancing the time cost and the material cost of the marketed goods that are used as inputs in the household production process. Becker [1960] also provides another view in which parents decide the number of children and their quality balancing the utility trade-offs between these decisions and the decision to consume an aggregate marketed good. In this line of research, children are treated analogous to consumption good.

There are several alternative theories of fertility choices in which children are treated as poor man's capital contrast to the consumption good view of the above literature. One such theory, introduced to the economic demographic literature by Cain [1981, 1983], focuses on the insurance aspect of childbearing. In this approach, the fertility decisions are guided by a lexicographic safety first preference ordering which was originally introduced in the agriculture literature. According to this principle, safety against old-age disability risks is of paramount concern in child bearing until certain number of children are born to ensure adequate hedging against such risks.[2] Another set of models formulate the old-age security motive for children within the life-cycle framework by assuming that parents receive certain fixed amount of transfers from children, the amount being determined by social norms (Nehar [1971], Raut [1985, 1990], and Willis [1980]). All these models of old-age security predict that the effect of infant mortality on fertility is either null or negative, in contrast to the predictions of the first approach. We will empirically reevaluate this effect estimating accelerated hazard models which have never been used in the literature for this purpose.

All these theoretical models are regarding the desired (completed) family size and ignore the effects of life-cycle events. Fertility decisions are, however, sequential in nature and interact with the evolution of a couple's various socio-economic characteristics over their life-cycles; some of the important life-cycle characteristics include husband and wife's education, earnings profile of the couple. An early attempt to model sequential fertility decisions was by Leibenstein [1957]. His concern was to model the higher order birth and he ignored the determinants of the first one or two children. He expressed all costs and benefits of having children in terms of utility and disutility. He assumed that a higher order birth is wanted for three types of utility - 1) consumption utility, 2) work or income utility, and 3) old-age security utility. The disutility from a higher order birth is due to 1) direct child rearing cost, and 2) other indirect costs such as income earning opportunities foregone by parents in raising the child. The couple will decide to have a higher order birth if the benefits exceed costs. While his theory allows fertility decisions to respond to changes in the life cycle events, by assuming that fertility could be controlled deterministically, his theory has no bearing on the birth intervals. Moreover the problem is not set up in a choice theoretic framework by explicitly specifying the utility function and the constraints.

Heckman and Willis [1975] improve upon these deficiencies. They formulated the hazard or risk of birth as the choice variable and studies how they evolve over the life-cycle of a couple as a function of the cost of contraception, birth parity, the time profile of earnings, and the cost of children in a discrete time dynamic programming framework. Newman and McCulloch followed a similar approach in continuous time to study the waiting time distributions. As it turns our these life cycle optimization problems are either impossible to yield an explicit optimal decision rule or it is computationally formidable to estimate a decision rule even in highly simplified models (see for instance Wolpin [1984]). Thus, in my analysis I refrain from specifying a particular utility function and solving the utility maximization problem explicitly. Instead, I parameterize the optimal decision rules regarding the timing of marriage and timing

2. Reformulating it as a switching regression model, Jensen [1990] finds support for old-age security motive using the Malaysian Life Survey Data.

and spacing of birth as follows:[3]

Consider only the family formation decisions of the households, and assume all other decisions such as savings and labor supply to be exogenously given. The only family formation decisions that we are concerned with in this paper are timing of marriage and timing and spacing of births. Typically a woman will visit the following biological states birth, pregnancy resulting in a live-birth, and inability to conceive as a result of menopause or death of husband. We restrict our analysis to the subset of states, S = {marriage, pregnancy leading to a live-birth, inability to conceive} only. No economic agent has full control over any of these timings, although they can have partial control by their choice of a mix of instruments. For instance, using such instruments as dowry, health and beauty care, efforts on scholastic performance, and establishing social connections, a woman can partially control her timing of marriage. Similarly, using a mix of contraceptive methods such as complete abstinence, pills, abortions, breast feeding and coital frequency a woman can partially control her timings of births.

Let T be the duration of an event such as timing of marriage, duration between marriage and the first live-birth, or between the first and the second live-births etc. These events will be referred to as event0, event1, and event2 respectively. Let us first talk about the determinants of birth intervals. Let u be a family planning strategy such as a sequence of contraceptive methods that the couple may adopt to control the duration of the event. Let U be the set of all feasible strategies. Such biological endowments as fecundability of the couple, which the couple might have some knowledge[4] about but is an unobservable to the econometricians, will also affect the duration of the event. Similarly for event0, the type of controls are search intensity of finding a partner, and the biological endowments include physical appearance. We will lump all biological endowments that affect an event into one real variable η. Hazard rate of an event T is defined as follows :

$h(t|\ u, \eta) \equiv$ probability that the event T occurs in the time interval $(t, t+dt)$ given that it has not occurred until t and given the family planning strategy u is used and the value of individual specific unobserved heterogeneity is η. (1)

Let us assume that $u = 0$ represents no family planning and η is standardized to mean zero in the population so that $\eta = 0$ represents a woman with average level of biological endowments. A baseline hazard function is defined as $\lambda_0(t) \equiv h(t|\ u = 0, \eta = 0)$.

3. One important reason for specifying a formal sequential optimization problem is to obtain parameter restrictions implied by the theory and then to incorporate these restrictions in the estimation procedure to obtain more efficient estimates of the parameters and to test the theory. Since we know that we cannot derive such restrictions without assuming very simple functional forms, I will rather keep generality of the functional forms and parameterize the optimal decision rules in stead of deriving them from restrictive functional forms for utility and hazard functions.

4. The couple might not know about their fecundability to start with and may learn over time between marriage and first child.

In our terminology, Heckman and Willis assumed that the effect of a family planning strategy or biological endowments is to scale the base line hazard function proportionately up or down as follows;[5]

$$h(t \mid u, \eta) = \lambda_0(t) \cdot \psi(u, \eta), \quad \text{where} \quad \psi(u, \eta) > 0 \tag{2}$$

Let X denote the covariates whose values represent the information available to our couple at time t. The mix of partial control that the couple adopt during the period of the event T will depend on the objective function and the accumulated information at a given time during their life-cycles. For instance, if the couple have not been able to get alternative source of old-age support such as sufficient stock of assets and pension funds, they might like to have more children and hence adopt less contraception and shorter durations. For another example, suppose the couple have preferences for sons. Once they have certain number of sons they might use more contraception leading to longer duration. Thus changes in these life-cycle events affect a couple's contraception behavior and hence the durations between births. Biological endowments of the couple will also affect their optimal contraception strategy. Thus, under quite general conditions, an utility maximizing optimal contraceptive strategy will be, $u = u(X, \eta)$, a function of X and η. If we parameterize the composite function $\psi(u,(X,\eta),\eta) = e^{x'\beta+\eta}$, we get the proportional hazard model for the observed timing of marriage and timing and spacing of births for our couple:

$$h(t \mid X, \eta) = \lambda_0(t)\, e^{x'\beta} \tag{3}$$

Proportional hazard models for timing of marriage and timing and spacing of births could be misleading description of observed choices and may yield faulty policy prescriptions for reasons explained as follows:

Suppose the base-line hazard function is parameterized as loglogistic:[6]

$$\lambda_0(t) = \frac{\theta\left(\frac{t}{\alpha}\right)^{\theta-1}}{\alpha\left(1+\left(\frac{t}{\alpha}\right)^{\theta}\right)}, \quad 0 < \theta, \alpha < \infty \tag{4}$$

To illustrate our point simply, assume that there is no heterogeneity in the biological endowments and that we have only one convariate X say the woman's employment status, and it takes two values: $X = 0$ (unemployed) as $X = 1$ (employed). In a large random sample, we expect to observe that an unemployed woman will most likely not control her natural timing of birth and her hazard rate for a birth will look similar

5. To be more precise, Heckman and Willis assumed that the couple choose the hazard rate itself not the controls u. They further assumed that the optimally chosen hazard functions belong to the proportional class. It is easy to see this is equivalent to our reinterpretation of their work.

6. Equation (4) is an appropriate model in our context since it can generate the right kinds of shapes for the baseline hazard rate. For instance, if $\theta \le 1$, the $\lambda_0(t)$ is a monotonically decreasing function of t (this is appropriate for the event age at marriage); if $\theta > 1$, $\lambda_0(t) = 0$ at $t = 0$, monotonically increases to a maximum at some $t > 0$ and then it monotonically decreases to zero as $t \to \infty$ (this is appropriate for the birth intervals).

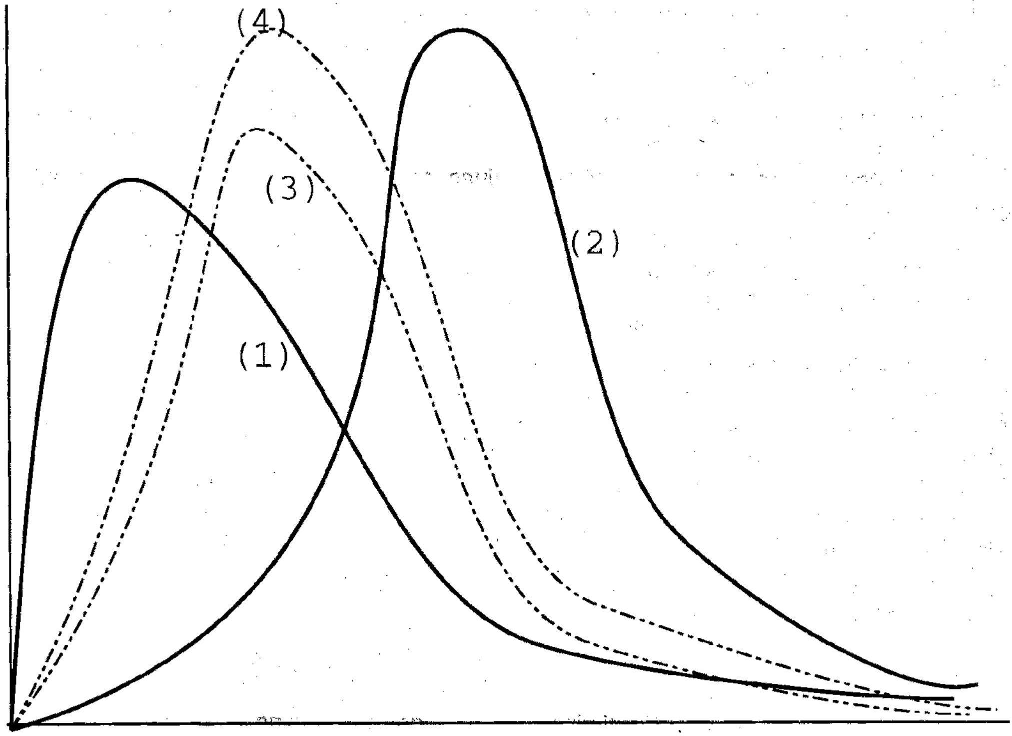


Figure 1 : Loglogistic proportional hazard function, $h(t \mid X, \eta) = \lambda_0(t)\, e^{X'\beta}$ fitted on data generated by accelerated hazard model.

to curve (1) in figure 1; on the other hand, the employed woman's optimal choice most likely will be to have a child with very low probability up to some point of time and thereafter to accelerate the chances of a birth as fast as possible.[7] The hazard functions which have such kind of acceleration property are depicted in curves (1) and (2) in figure 1 are known as *accelerated hazard model,* which we describe in details later.

Economists generally use proportional hazard models for economic duration data which impose severe restrictions. The proportionality assumption implies that the ratio $h(t_0 \mid X, \eta)/h(t_1 \mid X, \eta)$ depends only on t_0 and t_1 and not on X and η, i.e., the ratio is the same for both an employed woman and an unemployed woman; which have the further implications that the employed woman's hazard function will be an upward proportional shift of the hazard function of the unemployed woman. If we fitted a proportional hazard accelerated model on data generated by hazard functions (1) and (2), the estimated proportional hazard functions of the unemployed

7. It is not difficult to find a strategy in U that will allow her to have such a desired timing of birth.

and employed woman will look respectively like the curves (3) and (4) in figure 1. Thus we may come to the wrong inference that the effect of employment is to make a birth interval shorter.

Unlike in the proportional hazard models where the parameterization is done for hazard function, in accelerated hazard framework parameterization is done at a different level. Let the random variable T_0 denote the duration of an event conditional on the values of all covariates equal to zero; or in other words, T_0 could be viewed as the duration of the event for a woman drawn randomly from the population of women who do not control their fertility behaviour. The harzard rate of T_0 is given by $\lambda_0(t)$. In this sense, the hazard function $\lambda_0(t)$ represents the natural hazard rate and T_0 denotes the natural or base-line duration of the event. The effect of a covariate $X = 1$ in this model is to scale the natural duration, T_0, up or down by e^{β} according as β is positive or negative. Thus, while in proportional hazard framework the effect of such a regressor is to move the beseline hazard rate proportionately up or down, in accelerated hazard framework the corresponding effect is to move the base line duration T_0 proportionately up or down. More formally an accelerated hazard model specifies:

$$T = e^{X'\beta + \eta} \, T_0 \tag{5}$$

or equivalently,

$$\log T = \eta + X'\beta + \sigma e \tag{6}$$

where $T_0 \equiv e^{\sigma e}$, and $\sigma > 0$ is a scale parameter, e has a distribution independent of X, and β is the set of parameters of interest. The distribution of e can be derived from the baseline hazard distribution $\lambda_0(t)$ of T_0. To formulate the likelihood of a sample we need to derive survival function and the hazard function of T conditional on the values of X and η:

Let $S(t|X)$ be the survivor function, that is the probability that the event has not occurred until t, then we have

$$\begin{aligned} S(t|X, \eta) &= Prob\{T > t|X, \eta\} \\ &= Prob\{T_0 > te^{-X'\beta-\eta}\} \\ &= S_0(te^{-X'\beta-\eta}) \end{aligned} \tag{7}$$

where $S_0(t)$ is the survival function corresponding to the base-line hazard function $\lambda_0(t)$. Now we can derive the hazard function of T from $S(t|X, \eta)$ as follows:

$$h(t \mid X, \eta) = -\frac{d}{dt} \log S(t \mid X, \eta)$$

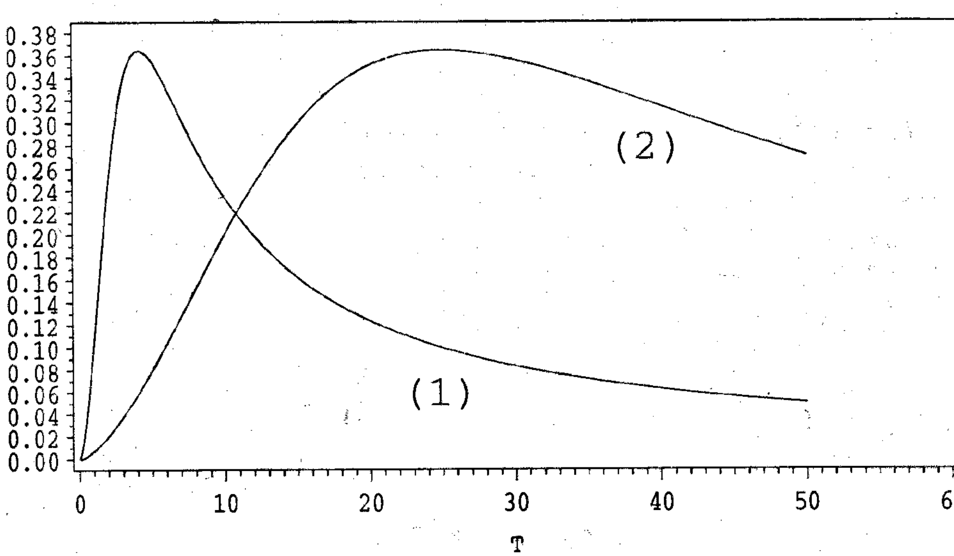
$$= \lambda_0 \left(te^{-x'\beta - \eta} \right) e^{-X'\beta - \eta} \tag{8}$$

For the previous example with one covariate, no unobserved heterogeneity and loglogistic base-line hazard function, the plot of the hazard functions corresponding to $X = 0$ and $X = 1$ are shown in figure 2. Notice the close resemblance of the curves (1) and (2) of the loglogistic accelerated hazard function in figure 2 with the curves of the observed hazard functions (1) and (2) in figure 1.

3.ECONOMETRICSPECIFICATION

It is easy to see that if we assume the regressors to be time varying, the specification in (6) will become a non-linear regression equation, the form of which will depend on a particular specification of $\lambda_0(t)$. However, the econometric techniques

Graph of loglogistic Accelerated Hazard Rates
(1) corresponds to X = 0
(2) corresponds to X = 1



Parameter values are: ($\alpha = 3.5$, $\theta = 2.5$, $\beta = 1.8$)

Figure 2 : Loglogistic accelerated hazard function : $h(t \mid X, \eta) = \lambda 0 \left(te^{-X'\beta} \right) e^{-X'\beta}$.

to handle time varying covariates in accelerated hazard framework are not yet developed in the literature, so we assume in this paper that the covariates are not time varying within the duration of an event.

The advantage of treating (6) as an accelerated hazard model than a linear regression model is that the former can use information from censored durations whereas the latter throws them out. The likelihood of the sample could be calculated as follows: Suppose we have data of the following type:

$$\{ t_j , \delta_j X_j : j = 1 , 2 , \ldots n \}$$

where t_j is either the completed duration or the duration of the event until the survey date, and $\delta_j = 1$ if the jth observation is censored, otherwise $\delta_j = 0$. The likelihood of the sample is given by

$$L (\beta , \eta_1 \ldots \eta_n) \equiv \prod_{j=1}^{n} (S (t_j | X, \eta_j))^{1-\delta_j}, h (t_j | X, \eta_j) (S (t_j | X, \eta_j))^{\delta j} \tag{9}$$

where $(\eta_1 \ldots \eta_n)$ is a vector of incidental parameters or fixed effects corresponding to the values of the unobserved heterogeneity and left out variable. We maximize the above likelihood for each event separately. Three important econometric issues in this connection are: First, is the accelerated hazard framework more appropriate for modeling economic duration data than the commonly used proportional hazard framework? Second, how to control for unobserved heterogeneity if it has significant effects on the parameter estimates? Third, how to choose between models?

Sueyoshi [1991] among others developed tests for proportionality assumption of the hazard rate and finds evidence against it in employment duration data for the U.S. We do not carry out any specification testing here, but we estimate a Weibull model (which belongs to both accelerated hazard class and proportional hazard class) and a loglogistic accelerated hazard model and compare the sensitivity of the parameter estimates of these two models from two different classes. To choose between non-nestd models, an appropriate approach would be to carry out Cox's non-nested specification testing and Pearsonian goodness of fit testing procedures, but there are no standard statistical packages that can perform these tests readily for duration models. As a first step we adopt a simpler strategy as follows: We estimate Weibull and loglogistic models for the baseline hazard function ignoring the unobserved heterogeneity. First we check if the parameter estimates of the regression coefficients vary "substantially" for these two models; if we find substantial variation then we use the Akaike Information Criterion (which essentially compares the maximized likelihood value of the sample under two models) to choose between these two models.

4. DATA SET

We use the 1976 Malaysian Family Life Survey data for our analysis. This data set contains the event history data on 1262 households drawn randomly from private households consisting of at least one married woman of age less than fifty. These households represent quite well all the socio-economic strata in the country, and the data set has passed many consistency checks (On data reliability, see Haaga [1981]). Malaysian population consists of three racial groups - Malay, Chinese, and Indians. Malays make up about 50% of the total population. The following are the variables that we use in this study:

In the economic demographic literature, such observed heterogeneity as MISCRG

Choice of variables :

ALT	=	1 if WEALTH > 0, and = 0 otherwise
C-AGE	=	age in months at last effective live birth divided by 100
CHLDTH	=	number of children died after age of six months
DOWRY	=	amount of dowry paid during marriage divided by 100
ED-LEVEL	=	level of education of the mother in number of years divided by 10
HEARNG	=	husband's monthly earnings divided by 1000
INFNTDTH	=	number of children died before age of six months
MISCRG	=	number of miscarriages up to the present time
MON_SEP	=	effective number of months the couples were geographically separated, divided by 10
RACE1	=	1 if the household is Malay, and = 0 otherwise
RACE2	=	1 if the household is Chinese, and = 0 otherwise
RURAL	=	1 if the household is in the rural sector, and = 0 otherwise
WEARNG	=	wife's monthly earnings divided by 1000
WEALTH	=	value of wealth divided by 1000
N_SON	=	number of surviving sons
OLDAGE	=	1 if the couple expect old-age support from their children and
	=	0 otherwise

and MON_SEP are generally not controlled for, but we want to control for these. The OLDAGE variable is an attitudinal variable and recorded only during the survey period; so the variable can have severe measurement errors when projected to the earlier years of a couple's life-cycle. The wealth is defined as the total value land, houses and building that are owned by the household.

Table 1 provide the summary statistics of these variables separately for each event.

5. EMPIRICAL RESULTS

We first examine the sensitivity of the parameter estimates to model specifications. In tables 2 and 3 respectively the estimates from Weibull and loglogistic models are reported. We include all variables of interest in these models. To be consistent with the computer print-outs, we relabel the events as follows: event0 = marriage, event1 = $0 \rightarrow 1$, event2 = $1 \rightarrow 2$, and so on. Comparing the parameter estimates from these two tables we find that most of them agree in sign and significance and a few disagree. For instance, the estimates of the effects of education level for the events $1 \rightarrow 2$ and $2 \rightarrow 3$ are significantly negative in Weibull model but they are not significant in loglogistic model. Similarly, the effect of husband's earnings is significantly positive for Weibull model but not significant for loglogistic model. The only other

Table 1
Summary statistics for selected variables

		Marrg	0 → 1	1 → 2	2 → 3	3 → 4	4 → 5	5 → 6	6 → 7	7 → 8
ALT	MEAN	0.351	0.430	0.497	0.537	0.551	0.588	0.643	0.678	0.635
	STD	0.478	0.495	0.500	0.499	0.498	0.493	0.479	0.468	0.482
CHLDTH	MEAN	--	--	0.082	0.122	0.184	0.254	0.342	0.458	0.470
	STD	--	--	0.280	0.363	0.440	0.536	0.626	0.725	0.723
C_AGE	MEAN	226.626	232.573	250.009	276.613	300.490	322.678	347.062	368.452	387.484
	STD	48.601	52.206	50.113	51.294	50.362	49.422	47.137	47.881	46.993
ED_LEVEL	MEAN	3.810	3.858	3.470	3.114	2.774	2.572	1.881	1.477	1.904
	STD	4.754	4.724	4.589	4.652	4.700	4.979	4.937	4.866	7.827
HEARNG2	MEAN	2.157	1.674	1.679	1.876	1.827	1.677	1.602	1.333	1.533
	STD	29.042	3.424	3,534	4.012	3.388	2.572	2.303	1.642	2.629
INFNTDTH	MEAN	0.000	0.056	0.107	0.148	0.199	0.243	0.291	0.272	0.278
	STD	0.000	0.246	0.362	0.409	0.486	0.540	0.572	0.590	0.579
MISCRG	MEAN	0.000	0.366	0.397	0.471	0.550	0.549	0.851	0.686	0.713
	STD	0.000	0.754	0.854	0.894	0.983	0.942	1.190	1.091	1.254
MON_SEPE	MEAN	--	1.970	3.404	3.210	4.669	4.355	3.245	2.847	4.296
	STD	---	10.646	13.552	12.957	16.467	15.152	12.938	12.945	16.192
RACE1	MEAN	0.482	0.488	0.480	0.469	0.445	0.454	0.538	0.550	0.499
	STD	0.500	0.500	0.500	0.499	0.497	0.498	0.499	0.498	0.501
RACE2	MEAN	0.390	0.384	0.395	0.402	0.420	0.393	0.334	0.305	0.351
	STD	0.488	0.486	0.489	0.491	0.494	0.489	0.472	0.461	0.478

.....Table 1 continues in the next page

Table 1 continues ...

		Marrg	$0 \rightarrow 1$	$1 \rightarrow 2$	$2 \rightarrow 3$	$3 \rightarrow 4$	$4 \rightarrow 5$	$5 \rightarrow 6$	$6 \rightarrow 7$	$7 \rightarrow 8$
RURAL	MEAN	0.588	0.583	0.583	0.601	0.621	0.637	0.679	0.709	0.693
	STD	0.492	0.493	0.493	0.490	0.485	0.481	0.467	0.455	0.462
WEARNG2	MEAN	0.424	1.129	0.456	0.407	0.405	0.396	0.348	0.335	0.344
	STD	1.296	21.295	1.217	0.757	0.772	0.741	0.576	0.576	0.552
WEALTH	MEAN	1382.31	5341.36	11246.4	16520.4	19058.6	30843.0	18029.9	53406.0	9308.81
	STD	29019.2	66686.6	122889	155265	182588	288623	261185	540678	25955.6
N-SON	MEAN	0.000	0.000	0.990	1.542	2.034	2.461	2.981	3.521	3.922
	STD	0.000	0.000	0.709	0.851	1.019	1.128	1.306	1.414	1.580
Number of observaton		1198.00	1418.00	1140.00	967.000	789.000	614.000	530.000	522.000	345.000

serious disagreement is for the effect of infant death for the event $2 \rightarrow 3$, which is significant in loglogistic but not in Weibull. All other parameters have similar estimates in both models. This calls for statistical analysis to investigate whether the disagreements are due to the proportionality assumption or due to not controlling for unobserved heterogeneity. I suspect that sensitivity might be due to the proportionality assumption since previous research has shown that within the proportional hazard class the parameter estimates of the covariates are not "very" sensitive to the specification of the base line hazard function when unobserved heterogeneity is ignored. (see Trussel and Richard [1985], and Raut [1989], for instance).

Notice that loglogistic model has uniformly higher likelihood values than the Weibull model. Thus using the Akaike Information Criterion we find evidence for our contention that the accelerated hazard framework is more appropriate for modeling economic duration data than the proportionality hazard framework. Using this fact that the disagreements of parameter estimates are not widespread, we draw inference about our hypotheses on the basis of the estimates from the loglogistic accelerated hazard model.

5.1 *Replacement effect*

Replacement effect measures the responsiveness of the fertility decisions to an infant or child death. In the literature there has been some dispute as to whether infant/child mortality is exogenous or it depends on the number of children in the

Table 2
Parameter Estimates from loglogistic model

	Marrg	0→1	1→2	2→3	3→4	4→5	5→6	6→7	7→8
Intercept	5.289	3.298	2.820	2.447	2.301	2.679	2.335	2.550	2.348
t-ratio	327.2	24.301	27.275	19.284	14.086	14.458	10.813	12.344	7.439
Alternate source of oldage support	0.032	0.083	0.051	0.106	0.037	0.100	0.164	0.234	0.166
t-ratio	3.059	1.761	1.622	2.892	0.830	2.076	3.038	4.433	2.186
child death	--	--	-0.030	-0.072	-0.088	-0.082	0.021	0.005	0.027
t-ratio	--	--	-0.526	-1.463	-1.781	-1.838	0.482	0.159	0.558
Age at which the event started	--	-0.004	0.001	0.001	0.002	0.002	0.002	0.001	0.002
t-ratio	--	-6.677	1.589	3.193	3.875	2.606	2.736	1.491	2.581
Education level of the woman	0.014	-0.004	-0.003	-0.004	0.001	0.004	-0.003	0.003	-0.004
t-ratio	9.590	-0.786	-0.833	-1.052	0.211	0.866	-0.730	0.633	-1.155
number of sons	--	--	0.023	0.039	0.040	-0.008	0.019	-0.012	-0.051
t-ratio	--	--	1.049	1.872	1.854	-0.393	0.979	-0.708	-2.330
RURAL	-0.024	-0.027	0.033	0.003	0.126	-0.017	0.059	0.063	-0.024
t-ratio	-2.312	-0.556	0.959	0.084	2.595	-0.327	1.011	1.100	-0.305
Husband's earnings	-0.000	0.001	0.014	0.006	0.016	-0.009	0.010	-0.019	0.044
t-ratio	-1.127	0.099	2.162	1.045	2.060	-0.796	0.741	-1.247	1.918
Wife's earnings	0.042	0.006	-0.002	0.019	-0.016	0.009	-0.006	0.003	0.030
t-ratio	5.935	0.484	-0.148	0.709	-0.453	0.257	-0.126	0.062	0.427
Maximized loglikelihood	456.90	-1532	-745.7	-691.9	-608.6	-450.4	-354.6	-363.1	-274.5

* Other variables are omitted

Table 3
Parameter Estimates from Weibull model

	Marrg	0 → 1	1 → 2	2 → 3	3 → 4	4 → 5	5 → 6	6 → 7	7 → 8
Intercept	5.352	3.409	2.964	2.700	2.412	3.019	2.472	3.135	2.806
t-ratio	268.19	26.110	25.333	18.375	14.112	15.981	11.589	15.322	7.687
Alternate source of oldage support	0.085	0.019	-0.018	0.056	0.003	0.139	0.189	0.229	0.192
t-ratio	6.942	0.406	-0.487	1.299	0.068	2.892	3.631	4.233	2.416
child death	--	--	-0.333	0.015	-0.083	-0.036	0.034	0.002	0.056
t-ratio	--	--	-0.537	0.250	-1.725	-0.786	0.895	0.047	1.125
Age at which the event started	--	-0.003	0.001	0.001	0.003	0.002	0.002	-0.000	0.002
t-ratio	--	-5.932	3.147	2.410	4.300	2.564	2.915	-0.593	1.737
Education level of the woman	0.010	-0.011	-0.008	-0.010	0.001	0.004	-0.007	0.001	-0.007
t-ratio	5.789	-2.827	-2.632	-3.312	0.252	0.661	-1.648	0.149	-1.452
number of sons	--	--	-0.000	0.066	0.042	-0.019	0.034	-0.012	-0.060
t-ratio	--	--	-0.011	2.698	1.905	-0.859	1.711	-0.676	-2.673
RURAL	-0.003	0.050	0.063	0.032	0.144	-0.111	0.086	0.098	-0.035
t-ratio	-0.259	0.992	1.561	0.691	2.758	-2.027	1.473	1.680	-0.439
Husband's earnings	-0.000	0.001	0.011	0.008	0.021	0.028	0.031	-0.024	0.041
t-ratio	-1.062	0.071	1.908	1.269	2.581	1.902	2.595	-1.855	1.496
Wife's earnings	0.053	0.005	0.001	0.037	0.010	-0.010	-0.024	0.049	0.012
t-ratio	5.766	0.322	0.054	1.157	0.243	-0.232	-0.476	0.943	0.158
Maximized loglikelihood	229.06	-1565	-883.5	-798.2	-649.9	-478.7	-355.5	-368.7	-287.5

* Other variables are omitted

household. More children in a family may cause a higher rate of infant/child mortality because more members have to share the limited resources (see Heer [1983] on the controversy). Since in our hazard rate approach, we estimate the increment in the probability of having a child when there is an infant/child death for each parity separately, our estimates do not suffer from this bias.

As regards the replacement effect, notice that the estimates of child death is significantly negative only for the events $3 \rightarrow 4$ and $4 \rightarrow 5$ and the estimates for other events are not significant.[8]

It is quite likely that a child death will have not much effect on the duration of the first two to three births, since most parents tend to have their first one or two children soon after marriage and thus it is more likely that there may not be any child death during the first three birth intervals. But for higher order birth intervals, there might be higher incidence of child deaths. This is also clear from the mean child death figures for different parities in table 1. Notice that the parameter estimates for CHLDTH is significantly negative for parities as high as 4. This means that a significant proportion of parents in our samples want to have about 3 to 4 children. When they have a child death, they hurry up to have another child.

We also estimated the models for Malays and Chinese samples separately. The estimates of the loglogistic model are shown in tables 4 and 5 for Malays and Chinese respectively. It appears that the replacement effect is insignificant for Malays which is what Wolpin [1984] also reported. But for Chinese, the effect of child death is significantly negative not only for $2 \rightarrow 3$, and $3 \rightarrow 4$ but also for the higher order birth intervals $6 \rightarrow 7$ and $7 \rightarrow 8$. It appears then that Chinese have a higher desired family size than Malays. Thus larger family sizes for most Malays are not due to economic reasons but religious or some other reasons (Malays are in general muslims).

5.2 *Son preference hypothesis*

Notice that the estimated effect of number of sons strongly agree both in signs and significance in tables 2 and 3. Both tables show that the parents wait longer to have their second or third child if they already have a son; then the effect is insignificant until the seventh parity when the effect is just the opposite. Since for the higher order parities the estimates are either not significant or negative, the parents are having large number of children not because of non-preference but due to some other reason. The corresponding estimates for Malays and Chinese in tables 4 and 5 show that while the Chinese exhibits at most weak preference for sons, the Malays do not exhibit any such preference. Leung [1988] carrying his analysis on the Chinese sub-sample also finds evidence of weak preference for sons. Thus this result on the son preference hypothesis is robust with respect to the type of hazard models used by others.

8. We also included infant death as one of the regressors. The effect of infant death is positive whenever significant (i.e., $1 \rightarrow 2$, $2 \rightarrow 3$ and $5 \rightarrow 6$). This is not surprising since an infant death directly causes longer durations between two live-births (i.e., births where a child does not die as an infant). However, it has a possible negative effect if parents desire a large number of children. The net effect could be of either sign but more likely to be positive.

Table 4
Parameter Estimates from loglogistic model for the Malay population

	Marrg	$0 \to 1$	$1 \to 2$	$2 \to 3$	$3 \to 4$	$4 \to 5$	$5 \to 6$	$6 \to 7$	$7 \to 8$
Intercept	5.21	4.14	3.15	3.06	2.56	2.92	2.75	2.95	3.26
t-ratio	309.43	21.93	20.66	15.97	11.11	11.76	10.22	12.31	8.44
Alternate source of oldage support	0.01	0.18	0.05	0.09	0.03	0.07	0.13	0.51	0.09
t-ratio	0.73	2.27	0.90	1.63	0.47	1.03	1.78	2.33	0.97
child death	--	--	-0.11	-0.06	-0.07	-0.05	0.02	0.06	0.12
t-ratio	--	--	-1.51	-0.87	-1.17	-0.94	0.48	1.69	2.21
Age at which the event started	--	-0.01	0.00	0.00	0.00	0.00	0.00	0.00	0.00
t-ratio	--	-6.52	0.47	1.05	2.98	1.72	1.96	0.81	0.27
Education level of the woman	0.02	-0.02	-0.03	-0.04	-0.01	0.02	-0.01	-0.02	0.01
t-ratio	8.77	-1.64	-2.84	-3.19	-0.94	1.67	-0.61	-1.34	0.22
Number of sons	--	--	0.01	0.04	0.02	-0.03	0.00	0.01	-0.07
t-ratio	--	--	0.42	1.21	0.59	-1.00	0.17	0.33	-2.51
RURAL	0.01	-0.14	0.14	0.08	0.17	0.06	0.08	0.21	0.10
t-ratio	0.35	-1.65	2.33	1.09	2.09	0.80	0.92	3.11	0.94
Husband's earnings	0.02	-0.02	-0.02	0.02	0.01	-0.04	-0.07	-0.01	-0.00
t-ratio	3.71	-0.69	-0.87	0.80	0.26	-1.64	-2.55	-0.43	-0.18
Wife's earnings	0.04	0.02	-0.00	0.10	-0.02	-0.05	-0.06	-0.38	-0.07
t-ratio	3.01	1.11	-0.01	1.70	-0.34	-1.02	-0.74	-5.80	-0.72
Maximized loglikelihood	224.31	-783.3	-379.3	-333.8	-257.1	-175.5	-152.3	-150.1	-99.66

* Other variables are omitted

Table 5
Parameter Estimates from loglogistic model for the Chinese population

	Marrg	0 → 1	1 → 2	2 → 3	3 → 4	4 → 5	5 → 6	6 → 7	7 → 8
Intercept	5.45	2.70	2.74	2.47	2.22	2.78	1.91	2.56	1.79
t-ratio	383.82	13.75	17.62	12.04	8.13	8.54	4.55	5.43	2.67
Alternate source of oldage support	0.04	0.04	0.05	0.12	0.06	0.09	0.10	0.18	0.28
t-ratio	2.91	0.56	1.06	2.16	0.82	1.16	1.06	1.80	2.20
Child death	--	--	0.18	-0.15	-0.17	-0.07	0.00	-0.17	-0.19
t-ratio	--	--	1.44	-1.44	-1.74	-0.75	0.03	-2.21	-1.85
Age at which the event started	--	-0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00
t-ratio	--	-1.63	1.05	2.68	2.51	0.99	2.50	1.41	1.68
Education level of the woman	0.01	0.00	0.01	-0.00	0.01	0.00	-0.01	0.01	0.00
t-ratio	5.11	0.10	1.07	-0.17	0.71	0.19	-0.50	0.51	0.19
Number of sons	--	--	0.02	0.05	0.08	0.04	0.02	-0.08	-0.04
t-ratio	--	--	0.58	1.49	2.17	1.22	0.70	-2.27	-1.04
RURAL	-0.04	0.05	-0.05	-0.01	0.14	-0.07	0.11	0.11	-0.05
t-ratio	-2.55	0.70	-1.12	-0.09	1.99	-0.92	1.19	1.08	-0.37
Husband's earnings	-0.00	0.00	0.02	0.00	0.02	-0.01	0.04	0.00	0.16
t-ratio	-1.57	0.43	2.54	0.66	2.23	-0.82	2.13	0.09	2.70
Wife's earnings	0.03	-0.00	-0.01	0.02	0.01	0.03	0.06	0.15	0.09
t-ratio	2.97	-0.19	-0.38	0.57	0.24	0.51	0.88	2.07	0.88
Maximized loglikelihood	208.55	-542.1	-255.2	-260.2	-240.5	-167.3	-115.3	-107.5	-106.1

* Other variables are not reported

5.3 *Old-age security hypothesis*

One of the objectives in collecting the Malaysian Family Life Survey Data was to address the old-age security related issues. The data set does contain information on the attitudes of the parents as to whether they want to depend on their children for old-age support or not, and whether they expect pension from alternative sources. Unfortunately, this information pertains to the attitude of the parents only during the survey date and cannot be postulated as the attitude in the past years of their life-cycles. Attitude may change over time as a response to changes in the events over a couple's life-cycle. For instance, in the beginning of their reproductive period, a couple may expect not to depend on their children, hoping to get a job in the organized sector that provides old-age pension, or to accumulate enough assets. They may, however, later find that their expectations did not realize. As a result they may change their attitude towards children for old-age support. So it would be misleading to interpolate the current attitude to the past years of the couple's life-cycle. One way out of this problem could be to drop all events other than the event on the survey date and estimate the model on that restricted sample; but then we find that most women in the sample have censored durations and the parameter estimates based on this sample will obviously produce biased results.

An appropriate approach would be to find some suitable instruments for the alternative sources of old-age support other than children. The stock of wealth at each point in time is a more appropriate predictor of parents' expectations of the alternative source of old-age support. In our data set we have only recall information on the evolution of wealth over one's life-cycle. This information has a lot of variation and possibly high errors in recalling the exact amount. Therefore, we create a time varying dummy variable ALT as ALT = 1 if wealth is positive, otherwise ALT = 0. We also tried other cut-off points for defining ALT, but the estimates did not change much. The results are reported in tables 2-5. It is clear that the parents with alternative source of old-age support have longer duration between births. Here the results differ somewhat in the two models. The loglogistic estimates are significantly positive for all events except $3 \rightarrow 4$ and the effect being weakly significant for the first child. The estimates from Weibull model show the effect of ALT to be insignificant for the first three children. From both estimates it is fairly clear that the alternative source of old-age support have negative effects on the duration between births and hence on the number of children demanded.

We carried out another direct test of old-age security hypothesis by estimating a logit model on the sample restricted to the survey date as follows:

$$Prob\ \{\ \text{OLDAGE} = 1\ \} = \frac{e^{X\beta}}{1 + e^{X\beta}}$$

where X is a set of covariates such as RACE1, RACE2, AGE, EVENT (i.e., number of children), ALT, HEARNG. Originally we included many other variables such as wife's earnings, number of sons, but since they were not significant, we dropped those variables from our final logit model; the estimates of the final logit model are

Table 6
Parameter Estimates for logit model of old-age security hypothesis

Variables	Parameter estimates
Intercept	-4.661 (7.62)
Malay dummy	-1.103 (3.14)
Chinese dummy	0.515 (2.03)
Age of the mother	0.006 (4.99)
Number of children already born	0.042 (1.24)
Alternate source of oldage support	-0.378 (2.08)
Husband's earnings	-0.0002 (3.43)

t-ratios are in parentheses.

shown in table 6.

Here again we find that if parents have alternative source of old-age support, they have lower probability of depending on children for such support. Similarly, husband's income has significantly negative effect. Notice also that older the mother gets, higher is her chances of depending on children for old-age support. Another interesting fact that emerges from table 6 is that Malays (i.e., RACE1) has lower chances of depending on children for old-age support than the Chinese parents. The last fact is not surprising since it is well known that the Malaysian government gives a log of subsides and privileges to Malays than to Chinese and Indians.

5.4 *Other effects*

From tables 2 and 3 we notice that if the women have alternative source of old-age support, or higher education, or higher wage earnings or if she is Chinese, she marries later. Since we have choice based samples, namely we have information only on the married women with one child; some of the young women who are not yet married are omitted from the sample and thus will bias the effect of some of the covariates on age at marriage. Therefore, the estimates of the marriage event should be interpreted cautiously. We also find that while husband's earnings generally have positive effect on the first few birth intervals, wife's earnings has no effect.

Mother's education level has no significant effect on the birth intervals.

6. CONCLUSIONS

In this paper we have formulated the age at marriage and duration between live-births as accelerated hazard model and used the 1978 Malaysian Family Life Survey Data on about 1262 women to examine if there is empirical evidence for the replacement effect, old-age security hypothesis and sex-preference hypothesis. The overall sample contains all three races - Malays, Chinese and Indian. We have also examined sensitivity of the parameter estimates to the specification of base line hazard function as Weibull model (which belongs to both proportional and accelerated hazard classes), and as loglogistic accelerated hazard model. In both cases, we have ignored unobserved heterogeneity.

We find that quite a few parameter estimates differ significantly in these two models, some parameters even have opposite signs. We also find that for all events, according to Akaike Information Criterion, the loglogistic accelerated hazard model performs better that the Weibull model. This suggests that accelerated hazard models might be more suitable for modeling birth intervals and age at marriage. This, however, needs to be verified on other data sets as well as using other criteria such as specification testing and goodness of fit testing procedures.

As regards the replacement effect, we find that in the overall sample, there is evidence for replacement effect until the fourth child. When we analyzed data on Malays and Chinese separately, we find that replacement effect is not significant for Malays, but for Chinese it is significant up to the seventh child. Therefore, it appears that the Chinese have higher desired family size than Malays. Since we know that Malays get more subsidies and benefits from the government than the Chinese and Indians, the above differential in fertility behavior could be in response to such economic factors.

Regarding the son preference hypothesis, we find that in the overall sample parents wait longer to have their second or third child if they already have a son; for higher birth intervals the effect is not significant. The Malays, however, do not exhibit preference for sons at any parity. Since the effect is significant only for lower order parities, we can conclude that larger family sizes of the Malaysian household are not because of preference for sons, but due to some other factors.

We find strong evidence for old-age security hypothesis, especially among Chinese. We carried out another direct test of the hypothesis, namely, we estimated a logit model using the binary response data on the question: Would the respondent like to depend on their children for old-age support? The statistically significant estimates of the model suggest that the Chinese have higher likelihood of depending on their children for old-age support than Malays, and higher is the husband's earnings, it is less likely that they would depend on their children for such support. It is quite likely that the couples with higher income for husbands have better access to private pension and other assets that can support them at their oldage.

REFERENCES

Becker, Gary (1960), "An Economic Analysis of Fertility", in *"Demographic and Economic Change in Developed Countries"*, a conference of the Universities National Bureau for Economic Research, Princeton, Princeton University Press.

Becker, Gary (1965), "A Theory of Allocation of Time," *Economic Journal* vol. 75 (299) : 493-517.

Behrman, Jere R., Robert Pollak and Paul Taubman (1982), "Parental Preferences and Provision for Progeny", *J. Polit. Econ.* vol 90 (February), 52-73.

Ben-Porath, Y. and F. Welch (1976), "Do Sex Preference Really Matter?", *Quarterly J. Econ.* 9, 285-307.

Cain, Mead (1981), "and Insurance: Perspectives on and agrarian change in India and Bangladesh", *Population and Development Review,* 7(3) : 453-474.

Cain, Mead (1983), "Fertility as an Adjustment to", *Population and Development Review*, 9(4):688-702.

De Tray, D. (1984), "Son Preference in Pakistan: An Analysis of Intensions vs. Behavior", *Research in Population Economics*, 5, 185-200.

Entwisle B. and C.R. Winegarden (1981), "Fertility as a deteminant and consequence of government sponsored pension programs in LDCs", paper presented at the *Annual Meeting of Population Association of America*, Washington, DC.

Haaga John (1981), "Validity and Reliability of Retrospective Life History Data from the Malaysian Family Life Survey", paper presented at the *Annual Meeting of Population Association of America*, Washington, D.C.

Heckman, J. and R.J. Wills (1975), "Estimation of Stochastic Model of Reproduction: An Econometric Approach" in N.E. Terleckyj (ed.), *"Household Production and Consumption"*, Columbia University Press, New York.

Heer, David M. (1983), "Infant and Child Mortality and the Demand for Children", *Population Review,* January-December, 7-16.

Hohm, C.F., F.J. Galloway, C.G. Hanson, and D.A. Biner (1984), "A Reappraisal of the social security-fertility hypothesis: a bidirectional approach", (mimeo) Sen Diego University.

Hotz, V.J. and R.A. Miller (1988), "An Empirical Analysis of Life Cycle Fertility and Female Labor Supply", *Econometrica.*

Jensen, E.R. (1990), "An Econometric Analysis of the Old-age security Motive for Child Bearing", *Int. Econ. Rev.* 31(4):953-968.

Kelly, A.C., P. Cutright, and D. Hittle (1976), "Comment on Charles F. Hohm's Social Security and Fertility : An International Perspective", *Demography* 13 (November) : 581-586.

Leibenstein, H. (1957), *"Economic Backwardness and Economic Growth"*, New York, Wiley.

Leibenstein,H. (1974), "An Interpretation of The Economic Theory of Fertility : Promising Path or Blind Alley?", *J. Econ. Lit.* 12(2):457-479.

Leung Siu-Fai (1988), "On Tests for Sex Preferences", *J. Pop. Eco.* : 95-114.

Nehar, Philip A. (1977), "Peasants, Procreation, and Pensions," *Amer. Eco. Rev.*, Vol. 61(3) : 380-89.

Nerlove, M. and L.K,. Raut (1995), "Growth Models with Endogenous Population : A General Framework", (forthcoming) in (eds) M.R. Rosenzweig and O. Stark, *The Handbook of Population and Family Economics*, North-Holland.

Nerlove, M., A. Razin and E. Sadka (1987), *Household and Economy : Welfare Economics of Endogenous Fertility,* Academic Press, New York.

Newman, John and C.E. McCulloch (1984), "A Hazard Rate Approach to Timing of Births", *Econometrica,* Vol. 52(4).

Nugent J. (1985), "The Old-age Security motive for Fertility", *Pop, and Dev. Rev.* vol. 11(1) : 75-97.

Nugent J and T.I. Gillaspy (1983), "Old-age pensions and fertility in rural areas of less developed countries: Some evidence from Mexico", *Econ. Dev. and Cult. change,* vol 31(4): 809-829.

Olsen, R.J. and K.I. Wolpin (1983), "The Impact of Exogenous Child Mortality on Fertility : A Waiting Time Regression with Dynamic Regressors", *Econometrica,* 51(3); 731-749.

Raut. L.K. (1985), "*Capital Accumulation, Income Inequality and Endogenous Fertility is an Overlapping Generations General Equilibrium Model*", the second essay in unpublished Ph. D. dissertation, Yale University, New Haven.

Raut, L.K. (1989), "A Duration Analysis of Fertility Behavior of Malaysian Women : Partial Likelihood and nonparametric maximum likelihood estimates", Discussion Paper # 89-10, University of California, San Diego.

Raut. L.K. (1990), "Capital Accumulation, Income Distribution and Endogenous Fertility in an Overlapping Generations General Equilibrium Model", *J. Dev. Econ.* 34(1/2): 123-150.

Raut, L.K. (1992), "Effects of Social Security on Fertility and Saving : An Overlapping Generations Model", *Indian Economic Review,* vol. XXVII(1): 25-43.

Robinson, W.C. (1986), "High Fertility as risk insurance," *Population Studies,* vol. 40(2):289-298.

Sanchez, N. (1984), "Old-age security and Fertility : two new tests", (mimeo) Worcester, Mass : College of the Holy Cross.

Sueyoshi, G (1991), "Evaluating simple alternatives to proportional hazard model: Unemployment insurance receipt and the duration of unemployment", (mimeo), University of California-San Diego.

Trussel, James and Toni Richards (1985), "Correcting for unmeasured Heterogeneity in Hazard Models Using Heckman-Singer Procedure", (ed), Nancy B. Tuma, "*Sociological Methodology*", Jossey-Bass Publishers, Washington.

Wolpin, Kenneth I. (1984), "An Estimable Dynamic Stochastic Model of Fertility and Child Mortality", *Journal of Political Economy,* vol. 92 (51)