DISCUSSION PAPERS IN ECONOMICS

Working Paper No. 99-20

The Impact of Child Death on Birth Interval:
A Hazard Model Analysis of Micro Data from Nepal

Sameer Rajbhandary
Department of Economics, University of Colorado at Boulder
Boulder, Colorado

October 1999

Center for Economic Analysis
Department of Economics

University of Colorado at Boulder
Boulder, Colorado 80309

The Impact of Child Death on Birth Interval:
A Hazard Model Analysis of Micro Data from Nepal[1]

Sameer S. Rajbhandary[2]
Department of Economics, and Population Program, Institute of Behavioral Science

Abstract

This paper analyzes the effect of a child's death on the subsequent birth interval. First, using the life table method compared the cumulative fertility of an additional birth by survival status of the index child. Then, a logistic formulation of the discrete-time hazard model was used to estimate the risk of the event (closing of the birth interval by subsequent birth) or the hazard of next birth in any given month, starting from nine months after the birth of the index child. The model was extended to allow for time dependent effects of child death. Sets of biological and socioeconomic variables that are likely to affect fertility were also included in the model. The survival status of the index child was found to have a substantial effect on the risk of closing birth interval. The effect of child death was strong initially but declined over time. The empirical results also showed some evidence for the existence of son preference in Nepalese society.

JEL classification: J13

Key words: Index child, Birth interval, Discrete-time hazard model

[1] University of Colorado at Boulder, Department of Economics, Working Paper # 99-20
[2] I would like to thank Prof. Robert McNown, Prof. Jane Menken, and my colleague Nizam Khan.
Address for correspondence: Sameer S. Rajbhandary, Population Program, Institute of Behavioral Science, Campus Box 484, University of Colorado, CO 80309, USA. Fax: (303) 492 6924, e-mail: Sameer.Rajbhandary@colorado.edu

INTRODUCTION

There have been many studies on infant and child mortality, estimating the risk or hazard of dying at a certain age, and relating this to various sociodemographic and fertility variables such as previous and subsequent birth intervals. Comparatively few researchers have examined the impact of infant or child death on the subsequent birth interval. One such study was conducted recently by Gummer-Strawn, Stupp and Mei (1998) employing DHS (Demographic and Health Surveys) data from 46 developing countries (excluding Nepal). They found that mean birth intervals are 32 percent longer if the child survives than if the child dies in infancy. They also examined the time-varying effects of child death for five African countries and concluded that the effects for closing the birth interval are strongest within 24 months, drop off quickly, and after 30 months the occurrence of child death does not affect the propensity to close the interval. Another recent study by Park et al. (1998) found a strong effect of child death on subsequent fertility, with the effect becoming weaker with time

This paper uses the micro data from the *Nepal Fertility, Family Planning and Health Survey (NFHS) 1991* to examine the effect of a child's death on the subsequent birth interval. First, the cumulative fertility of an additional birth by survival status of the index child is compared using the life table method (Park et al.). Then, the logistic formulation of the discrete-time hazard model is used to estimate the risk of the event (closing of the birth interval by subsequent birth) or the hazard of next birth in any given month, starting from nine[1] months after the birth of the *index*[2] child. The relative risks when the index child has died compared to when the child is still alive are estimated to analyze the effect of the child death on the subsequent birth. The model is also extended to allow for time dependent effects of child death, that is, to see how the effect varies with the time since the birth of the index child. In other words, does it matter whether the index child is older or younger when it dies.

THEORETICAL FRAMEWORK

The effect of infant and child mortality on fertility operates through both physiological mechanisms and volitional behaviors. Preston (1978) identifies two mechanisms - *lactation interruption* effects and *behavioral replacement* - strategies that affect the subsequent birth interval in case of an infant or child death.

The child's death causes the cessation of breastfeeding, thus increasing the probability of return to ovulation, so also the conception of the next child, resulting in a shortened birth interval. The strength of this effect is found to be related to the intensity, frequency and duration of breastfeeding (Grummer-Strawn et al. 1998). This relationship is expected to be most influential in a society where breastfeeding is most common and is done for extended periods of time, and the use of contraceptives for birth spacing purposes is rare (Preston 1978).

Nepal can be characterized as one such society. Breastfeeding is nearly universal in Nepal; 97 percent of all children born in the period five years prior to the survey were breastfed. They were breastfed for an extended period with a mean duration of breastfeeding of 30 months. Only 28 percent of currently married women have ever used modern contraceptive methods. Permanent methods such as female and male sterilization accounted for 67 percent of those women practicing contraception (MOH [Nepal] 1993). This indicates that modern contraceptive methods have been used mostly for limiting births; its use for birth spacing is still very limited in Nepal.

The behavioral effect again might work in various ways. The death of one child may elicit a desire to have another one soon to replace the dead child, especially if it is a boy in a society with preferences for sons (Lehrer 1984). The behavioral response to this desire would be the cessation of contraception (possibly not very important in a low contraceptive prevalence country) and/or discontinuation of sexual abstinence (if the custom of abstinence from sexual relations after birth exists) (Van-Ginneken 1974). The behavioral mechanism might also work in the reverse direction; the unpleasant experience of child death may lead a couple to postpone having another child to avoid re-exposing itself to the risk (Preston 1978). These behavioral effects will be stronger in a society where women have control over their fertility and means to do so, as in high contraceptive prevalence countries. These effects are expected to be weaker in a low contraceptive prevalence country like Nepal.

Mainly, the following hypotheses will be investigated:

1. The death of a child raises the hazard of closing the subsequent birth interval and therefore shortens the interval.
2. The child death has a time-varying effect on subsequent birth interval. The effect declines over time, so that the effect is stronger if the index child is younger when it dies.
3. There is a preference for sons in Nepalese society.

a. The hazard of closing the subsequent birth interval is lower if the index child is male.
b. The effect of child death on the probability of subsequent birth is stronger if the index child who died is male.

DATA

This paper uses data from the *Nepal Fertility, Family Planning and Health Survey (NFHS) 1991* that contains information on fertility, family planning, infant and child mortality, and maternal and child health. Some information on the household characteristics was obtained from the household questionnaire, though its main objective was to identify women eligible for the individual interview. The questionnaire also included data on the timing of women's live births, the survival status of these births, and the interval to death of a child where applicable. For all births occurring in the five-year period prior to the survey (22,155 births), more detailed information including duration of breastfeeding, were also obtained.

The questionnaire was constructed with considerable help from the *Demographic and Health Surveys (DHS) Model B Questionnaire* (Ministry of Health [Nepal] 1993) which was designed for use in countries with low contraceptive prevalence (Pradhan et al. 1997). This survey was based on a multi-stage probability sample design: three stages of selection were involved for rural and two for urban samples and covered all 75 districts of the country. The survey was conducted over the seven-month period from August 1991 to February 1992. All ever-married woman in the age group 15-49 who had started living together with their husbands and who slept in the sample household the night before the day of interview were considered eligible for individual interview (de facto coverage). A total of 25,384 women from 24,754 households were interviewed. The sample comes from 1184 rural clusters and 62 urban clusters. A cluster consisted of a single *ward*[3] or more than a ward in case of the wards with less than 40 households. The households within the selected clusters are selected with the *probability proportionate to size* (PPS) using the random start and the sampling interval. The sample within the district was self-weighting. But the appropriate final set of weights incorporating both the design and non-response components have to be used for the national estimates.

This survey has suffered moderately from a systematic exclusion of women when compared with the 1991 census data (MOH [Nepal] 1993). This might have been caused by the misreporting of women's age in the household questionnaire in order to push them out of the eligible range and also by the misclassification of the women's age by the interviewers themselves in order to reduce their work load, as they conducted both the household and individual women interviews.

There is also a systematic displacement of children's birth dates further back in time especially for those born during the five year period prior to the survey (MOH [Nepal] 1993). A part of this might be a deliberate attempt by the interviewers themselves to avoid asking a large number of questions pertaining to those children. This is a common problem in such surveys (Dangol, Retherford and Thapa 1997).

The completeness of reporting of dates of child's birth of children, both alive or dead, and recently born or born many years before the survey, was found to be unusually high compared to similar data from past surveys. This may indicate that the interviewers were motivated to obtain the children's date of birth by any means as the answer code "don't know" was not acceptable (MOH [Nepal] 1993).

Though this survey also suffers from some of the usual limitations of retrospective surveys, it is believed to be of better quality than the earlier such surveys conducted in Nepal. The misreporting of the age of the child also appears to be better than the more recent survey, the 1996 Nepal Living Standards Survey (Dangol et al. 1997).

EMPIRICAL SPECIFICATION

1. Model

The discrete-time hazard model (Allison 1982, 1984; Trussell and Guinnane 1993) is formulated with the dependent variable, the risk of the birth interval being closed by subsequent birth, or the hazard, as a function of a set of independent variables or covariates. Hazard models allow both time varying covariates and *censored* observations.

The discrete-time hazard model written in logit form is:

$$\log[P_{it}/(1-P_{it})] = \alpha + \beta'\mathbf{X}_{it}$$

where,

P_{it}: the probability that an event occurs to individual i at time t, given that it has not already occurred and the individual i is still at risk

$\mathbf{X}_{it}$: a set of covariates

α, β: a vector of coefficients to be estimated

The equivalent logistic regression function is:

$$P_{it} = 1/[1 + \exp(-\alpha - \boldsymbol{\beta}'\mathbf{X}_{it})]$$

The likelihood function is:

$$\log L = \prod_{i=1}^{n} [\Pr(T_i = t_i)^{\partial_i} [\Pr(T_i > t_i)^{(1-\partial_i)}$$

Allison (1982:74-75) demonstrates that the log likelihood function is just the log likelihood for the regression analysis of dichotomous dependent variables. This implies that discrete time-hazard model can be estimated by using programs for the analysis of dichotomous data.

In the logistic formulation of the discrete-time hazard model, each discrete time unit for each individual is treated as a separate observation or unit of analysis. And for each of those observations the dependent variable take the value 1 if an event occurred to that individual at that unit of time, 0 otherwise. For example, if an event occurred at time 10, then 10 different observations would be created and the dependent variable take the value 0 for the first nine observations and 1 for the last or the tenth observation. The explanatory variables or the covariates take whatever value they had at that particular unit of time, so can easily accommodate time-varying covariates.

The creation of multiple observations for each birth interval is not problematic here, as it is not an adhoc method; rather follows directly from factoring the likelihood function for the data. However, the birth intervals contributed by the same woman tend to be more alike than two randomly chosen intervals. Pooling them together without considering the dependence can lead to biased estimates and standard errors (Allison 1995).

The estimated coefficients show the relationship between the covariates and the risk of the event (closing the birth interval) in any given discrete time unit. A positive or negative coefficient indicates that the risk of the event increases or decreases respectively as the value of the covariate increases. The interpretation is done similarly relative to the omitted or the reference category in the case of categorical or dummy variables.

2. Method

The logistic formulation of the hazard model is used in this paper. A woman becomes at risk of next birth, nine months after the birth of the index child (assuming nine months of gestation). So, the unit of observation is the month of exposure to the risk of next birth, starting from nine month after the birth of the index child. Each birth interval starting from the tenth month after the birth of the index child is represented by several observations, one for each month. In each observation, the dependent variable takes the value 1 if a birth occurred that month and 0 otherwise. For example, a birth interval of 20 months is represented by 11 observations starting from the tenth month, and for the first 10 observations, the dependent variable take the value 0 and it takes the value 1 for the last or the eleventh observation. The fixed covariates take the same value for each month, but the time-varying covariates may change with monthly observations. The estimated coefficients describe the relationship between the covariates and the risk of the event occurring in any given month. This is a more natural representation of the birth process as conception can only occur in the short time span of each month (Mroz and Meir 1990).

On the basis of previous work (Park et al. 1998; Grummer-Strawn et al. 1998; John et al.; Trussell et al.), the variables included in the model as likely covariates are: duration since the birth of the index child, survival status of the index child, parity and sex of the index child, the length of previous birth interval, age of the mother at the birth of the index child[4], education, current place of residence and religion of the mother. Some of the variables are categorical by nature, the others are treated as such to allow for nonlinearity. Duration (since the birth of the index child) variable is coded into 16 categories of three[5] months each starting from nine months to the maximum of 60 months. Parity is coded into three categories: First birth, Low parity (2 to 4), and High parity (5 or higher). Previous birth interval is coded into three categories: Short (less than 24 months), Medium (24 to 45 months) and Long (greater than 45 months). For the first birth, the previous interval is defined as the interval from the date of first union[6] to the first birth. The previous birth interval is included as the past event history of a woman to help control for some of the unobserved characteristics of a woman, such as her underlying fecundity, whose omission can bias the parameter estimates (Allison 1984). The age of the woman at the birth of the index child is coded into three categories: Less than 25, 25 to 35, and More than 35. These categories are chosen to reflect the commonly postulated decline in fecundity with age (Trussell et al. 1985). The education of the mother is also coded into three categories: None (never been to school), Primary (completed less than one to five grades), and Secondary (completed grade six or higher). The original coding in the data is retained for the religion variable. The survival status of the index child is treated as a time varying covariate with a lag of nine months (assuming gestation of nine months) to insure that the status of the index child only before the conception of the subsequent child is being considered. For example, death of the index child at age 11 months can only affect the birth of the next child at duration 20 months or greater. This is to avoid the possible reverse causality of a short subsequent birth interval causing the premature death of the index child (Gummer-Strawn et al. 1998). The duration variable is also time-varying: its value increases by one in each successive month of observation starting from the birth of the index child. The rest of the variables are fixed covariates that take the same value in every month of observation for the same index child or the same woman.

Only the live births are being considered here, as information is not available for others like still births etc. There were 22,155 live births during the five-year period before the survey. The births that occurred less than nine months prior to the survey were dropped, as the hazard of closing the birth interval is assumed to be zero in the first nine months of a child's life. Multiple births[7] and births with missing or inconsistent dates of births[8] were also excluded from the working sample. This resulted in final sample of 18360 births for which consistent information was available for all the variables included in the model. Only 6833 births were closed by a subsequent birth, and the rest were open intervals or *right-censored* (by the interview date). This represents 349936 (341186 weighted) months of exposure to the risk of birth.

Means and standard deviations of the variables used in the regression analysis are shown in Table 1[9]. At the birth of the index child, 42 percent of the women were younger than 25, 13 percent were older than 35, and the rest were between 25 and 35. Only 4.3 percent of the women live in urban areas, and 90 percent of the women have never been to school. The distribution of the index child across gender is quite even with 51.7

percent males. For the previous interval, about 28 percent and 23 percent lie in the categories, short (less than 24 months) and long (more than 45 months) respectively with the rest in the medium category (25 to 45 months). Forty nine percent of the births were low parity (2 to 4), 31 percent high parity (5 or higher), and the rest were first births.

3. Results

Life Table Analysis

The cumulative fertility of an additional birth by survival status of the index child is first compared using the life table method. The cumulative probability of an additional birth, 1-S(t) (S(t) being the "survival function" at time t), is shown in Table 2 for every three successive months from nine months since the birth of the index child. Within five years of birth, more than three-quarters of women had another birth. The probability was clearly higher if a child died than if it survived. At two years after the index birth, the probability was about 3 times higher for a woman whose child had died. The relative difference however declined over time and at four years after the index birth, it was only 29 percent higher.

The estimated hazard rate[10] of childbirth at the midpoint of each interval is shown in Figure 1. The rate first increased with time, reaching a maximum level of 0.043 at 37.5 months, and then began to decline. Figure 2 depicts the estimated hazard rate ratios for dead versus surviving index children (the category "died" being in the numerator). The ratio was initially very high, almost 7 at 13.5 months but declined with time reaching unity at 40.5 months after the birth of the index child.

These results are very similar to those found by Park et al. (1998) in the case of Bangladesh. This provides a background to more complete analysis that includes effects of covariates.

Regression Analysis

The regression results for the four different hazard models with the monthly probability of subsequent birth as the dependent variable are given in Tables 3.[11] Model 1 includes all the covariates listed in Table 1. The estimated coefficient on duration is statistically significant and positive for all duration categories compared to the reference category (9-<12) months after the birth of the index child. The coefficient for the category (36-<39) months is the largest indicating the maximum hazard during this time interval. The estimated hazard ratios for the covariate "duration since the birth of the index child" represented by 15 dummy variables representing 16 categories, is shown in Figure 3.

The estimated coefficients show that the monthly probability of next birth declines with parity. The odds are 20.1 percent ((exp(-0.2248)-1)*100), and 30 .1 percent lower for "low" (2 to 4) and "high" (5 or higher) parity births respectively, compared to the first birth.[12] The results also show positive association between the monthly probability of next birth and previous birth intervals. Compared to those with "medium" (25-45 months) prior interval, the odds of next birth is 8.9 percent higher for those with “short” (less than 24 months) interval and 14.3 percent lower for those with “long” (more than 45 months) interval.

The coefficients on the age dummy variables are all statistically significant and have the expected signs. The hazard is greatest for women who start the interval at ages less than 25 and smallest for those who begin at ages greater than 35. The odds are 29.5 percent higher and 40.6 percent lower respectively, compared to the reference group of women who start the interval at ages between 25 and 35. The relative risk (age less than 25/age greater than 35) is 2.18 ((exp(0.2588)/exp(-0.2512)).[13] Trussell et al. (1985) found the relative risk of 3.1, 4.6, and 3.4 in Philippines, Malaysia, and Indonesia respectively, in their study of birth interval lengths in those countries.

The risk also differs by place of residence, with the odds 15.7 percent lower for urban woman than for rural. The religion dummies used as a control for differences in cultural practices such as the timing and nature of the supplemental foods to be given to the child (which might affect fertility indirectly through its effect on breastfeeding) are all statistically significant. The odds are 8.5, 17.5 and 36.4 percent higher for Buddhist, Muslim and women of other religion comparatively, compared to Hindu women. Though the dummy variables for education of the mother are not statistically significant individually, they are jointly significant at 10 percent level as shown by the likelihood ratio test.

The odds are 15.2 percent lower if the index child is male than if it is female. The positive significant coefficient on the “dead” variable shows the expected higher risk of closing the birth interval if the index child is dead than if it is still alive. The odds are 2.3 times higher.

The joint likelihood ratio tests performed on variables also confirm the above results.

The interaction between the survival status of the index child with duration since the index birth are added in Model 2 to examine the time dependent effect of child death on the subsequent interval. The fit of the model improved substantially as indicated by the difference in the log likelihood and the likelihood ratio. The likelihood ratio test of this model versus the original one indicates that the improvement is statistically significant. The effect of child death was strong initially with the odds ratio of 9.94 at (9-<12) months after the index birth, then declined over time to 2.40, 1.52, and 1.37 at (21-<24), (33-<36), and (45-<48) months after the index birth respectively.[14]

The survival status of the index child and sex interaction is included in Model 3. The odds are 3.9 percent[15] higher if the index child who died is male than if the index child who died is female. This result along with the earlier result that the odds are 15.2 percent lower if the index child is male than if it is female shows some evidence of the existence of son preference. All the interaction terms are included in model 4.

The likelihood ratio test of the rest of the models versus Model 1 indicates significant improvement in each of those models with Model 4 being the most preferred one.

Conclusion

This paper demonstrates that the survival status of a child has a substantial effect on the risk or hazard of closing birth interval (or conception resulting in a live birth) and the effect is time dependent. The estimates show that the risk is 2.3 times higher if the index child is dead than if it is still alive, even after controlling for the variables that are likely to affect fertility. The effect of child death declines over time. The odds are 9.94, 2.40, 1.52, and 1.37 times higher at (9-<12), (21-<24), (33-<36), and (45-<48) months after the birth of the index child respectively. It matters whether the index child is older or younger when it dies. The odds are 15.2 percent lower if the index child is male than if it is female. However, the odds are 3.9 percent higher if the index child who died is male rather than if the index child who died is female. This indicates some evidence for the existence of son preference in Nepalese society.

Notes:

1. The hazard of closing the birth interval or the subsequent birth is assumed to be zero for the first nine months of a child's life, allowing for a nine month gestation for a subsequent birth.

2. The index child is defined as the child whose birth opens the birth interval under consideration.

3. Ward is the smallest political unit.

4. The age of the mother at the birth of the index child is calculated using the information on the date of birth of the child and the mother. The day of birth is randomly assigned (using random number generator) as only the month and the year are available. The age of the mother at the interview date is used if only the year of her birth is missing. In case the month of her birth is missing, the months are assigned randomly within the range (1 to 5, 6 to 12, or 1 to 12) depending on the information on the year of birth and the age at the interview date.

5. The exception is the last category (54 to $\leq$ 60) that is of six months duration. It is done so to avoid the problem in estimation due to very few cases.

6. The day of the first union is randomly assigned as only the month and the year of the first union are available. The age of the woman at the first union is used if only the year of the first union is missing. In case both the month and the year are missing, the months are randomly assigned.

The age at first union is *left-censored* at 12 if it is less than 12 for calculating the previous interval for the first birth.

7. Multiple births are dropped after calculating the previous and subsequent intervals for the singleton births.

8. The day of birth is randomly assigned as only the month and the year of birth are available.

9. The weights incorporate both the design and non response components which have to be used for the national estimates.

10. The estimated hazard rate is the change in log S(t) per unit time interval (Allison 1995).

11. The models were estimated using logistic regression procedure of STATA using the probability weights associated with each woman in the sample to compensate for design and non response components. Also the 'cluster' option is used with the variable 'womenid' (unique identification of the woman) which allows the relaxation of the assumption of independence across observations within clusters, that is, the same women in this case, and gives the correct standard errors. It also employs robust variance estimates, using White's procedure.

12. The estimation of the model with parity and its squared term indicated that the hazard peaks at around parity six.

13. Modeling with age and age squared instead showed that the maximum risk is at around the women's age of 21 years at the birth of the index child.

14. These odd ratios are calculated as follows:
odd ratio (dead vs alive) for each duration category = exp ($\beta_{dead}+\beta_{(dead*duration\ dummy)}$)
for the category (9-<12), odds ratio = exp(2.296723) = 9.9416
for (21-<24), odds ratio = exp(2.296723-1.422033) = 2.3981
for (33-<36), odds ratio = exp(2.296723-1.875908) = 1.5232
for (45-<48), odds ratio = exp(2.296723-1.985545) = 1.3650

15. The odds ratio for male index child dead vs female index child dead = exp(-0.1909753 + 0.2295933) = 1.0394

REFERENCES

Allison, Paul D. 1995. "*Survival Analysis using the SAS System: A Practical Guide.*" Cary, NC: SAS Institute Inc.

--------. 1984. *Event History Analysis: Regression for Longitudinal Event Data*. Beverly Hills, California: Sage Publications.

--------. 1982. "Discrete-Time Methods for the Analysis of Event Histories." Pp. 61-98 in *Sociological Methodology,* edited by S. Leinhardt. San Francisco: Jossey-Bass.

Dangol Bishnu Das, Robert D. Retherford and Shyam Thapa. 1997. "Declining Fertility in Nepal." *Asia-Pacific Population Journal* 12(1):33-54.

Grummer-Strawn, Laurence, Paul W. Stupp and Zuguo Mei. 1998. "Effect of a Child's Death on Birth Spacing: A Cross-National Analysis." Pp. 39-73 in *From Death to Birth: Mortality Decline and Reproductive Change*, edited by Mark R. Montgomery and Barney Cohen. Washington DC: National Academy Press.

John, A. Meredith, Jane A. Menken and A.K.M. Alauddin Chowdhury. 1987. "The Effects of Breastfeeding and Nutrition on Fecundability in Rural Bangladesh: A Hazards-Model Analysis." *Population Studies* 41(3):433-446.

Lehrer, Evelyn. 1984. "The Impact of Child Mortality on Spacing by Parity: A Cox-Regression Analysis." *Demography* 21(3): 323-337.

Ministry of Health [Nepal]. 1993. *Nepal Fertility, Family Planning and Health Survey 1991: Main Report*. Family Planning and Maternal Child Health Division, Planning Research and Evaluation Section. Kathmandu, Nepal.

Mroz, Thomas A. and David R. Weir. 1990. "Structural Change in Life Cycle Fertility during the Fertility Transition: France Before and After the Revolution of 1789." *Population Studies* 44(1):61-87.

Park, Chai Bin, Mohamed Ataharul Islam, Nitai Chakraborty and Andrew Kantner. 1998. "Partitioning the effect of infant and child death on subsequent fertility: An exploration in Bangladesh." *Population studies* 52:645-356.

Pradhan, Ajit, Ram Hari Aryal, Gokarna Regmi, Bharat Ban, and *Pavalavalli* Govindasamy. 1997. *Nepal Family Health Survey 1996.* Kathmandu, Nepal and Calverton, Maryland: Ministry of Health [Nepal], New ERA, and Macro International Inc.

Preston, Samuel H. 1978. "Introduction." Pp. 1-18 in *The Effects of Infant and Child Mortality on Fertility.* New York: Academic Press.

Trussell, James and Timothy Guinnane. 1993. "Techniques of Event-History Analysis." Pp. 181 -205 in *Old and New Methods in Historical Demography*, edited by David S. Reher and Roger Schofield. New York: Oxford University Press.

Trussell, James, Linda G. Martin, Robert Feldman, James A. Palmore, Mercedes Concepcion and Datin Noor Laily Bt. Dato'Abu Bakar. 1985. "Determinants of Birth-Interval Length in the Philipines, Malaysia, and Indonesia: A Hazard-Model Analysis." Demography 22(2):145-168.

Van Ginneken, Jeron K. 1974. “Prolonged Breastfeeding as a Birth Spacing Method.” *Studies in Family Planning* 5(6):201-206.

Table 1: Mean and Standard Deviations of the Variables

Variables	Mean	Mean (weighted)
Duration since index birth (months)		
9-<12	ref.	ref.
12-<15	0.1379	0.1373
	(0.3448)	(0.3442)
15-<18	0.1239	0.1235
	(0.3295)	(0.3290)
18-<21	0.1107	0.1105
	(0.3138)	(0.3135)
21-<24	0.0975	0.0974
	(0.2967)	(0.2965)
24-<27	0.0831	0.0829
	(0.2761)	(0.2758)
27-<30	0.0686	0.0685
	(0.2528)	(0.2525)
30-<33	0.0564	0.0565
	(0.2307)	(0.2308)
33-<36	0.0463	0.0466
	(0.2101)	(0.2108)
36-<39	0.0369	0.0372
	(0.1885)	(0.1893)
39-<42	0.0282	0.0283
	(0.1654)	(0.1659)
42-<45	0.0209	0.0210
	(0.1430)	(0.1434)
45-<48	0.0153	0.0155
	(0.1226)	(0.1235)
48-<51	0.0111	0.0114
	(0.1050)	(0.1060)
51-<54	0.0073	0.0075
	(0.0853)	(0.0865)
54-≤60	0.0057	0.0060
	(0.0756)	(0.0775)
Characteristics of the Index Child		
Survival status		
Alive	ref.	ref.
Dead	0.0564	0.0531
	(0.2306)	(0.2242)
Parity		
First birth	ref.	ref.
Low (2 to 4)	0.4940	0.5005
	(0.5000)	(0.5000)
High (5 or higher)	0.3051	0.2929
	(0.4605)	(0.4551)

Variables	Mean	Mean (weighted)
Previous interval (months)		
Short (less than 24)	0.2785	0.2785
	(0.4483)	(0.4482)
Medium (24 to 45)	ref.	ref.
Long (more than 45)	0.2293	0.2284
	(0.4204)	(0.4198)
Sex		
Male	0.5172	0.5164
	(0.4997)	(0.4997)
Female	ref.	ref.
Characteristics of the Mother		
Age at index birth (years)		
Less than 25	0.4192	0.4296
	(0.4934)	(0.4950)
25 to 35	ref.	ref.
More than 35	0.1328	0.1267
	(0.3393)	(0.3327)
Residence		
Urban	0.0427	0.0551
	(0.2022)	(0.2281)
Rural	ref.	ref.
Education		
None	ref.	ref.
Primary	0.0541	0.0574
	(0.2262)	(0.2327)
Secondary	0.0480	0.0586
	(0.2138)	(0.2350)
Religion		
Hindu	ref.	ref.
Buddhist	0.0834	0.0774
	(0.2765)	(0.2673)
Muslim	0.0310	0.0388
	(0.1733)	(0.1930)
Others	0.0154	0.0156
	(0.1230)	(0.1241)
Sample Size	349936	341186

Source: Derived from Nepal Fertility, Family Planning and Health Survey (NFHS) 1991.
Note: standard deviations are in parentheses.

Table 2: Cumulative Probability of Subsequent Birth

Months after the birth of the index child	Total	Index Child	
		Alive	Dead
0-<9	0.0000	0.0000	0.0000
9-<12	0.0053	0.0035	0.0228
12-<15	0.0268	0.0174	0.1135
15-<18	0.0540	0.0380	0.2006
18-<21	0.0867	0.0667	0.2679
21-<24	0.1392	0.1159	0.3466
24-<27	0.2186	0.1888	0.4785
27-<30	0.2883	0.2571	0.5571
30-<33	0.3556	0.3253	0.6145
33-<36	0.4207	0.3914	0.6693
36-<39	0.4912	0.4636	0.7237
39-<42	0.5484	0.5239	0.7548
42-<45	0.5964	0.5747	0.7795
45-<48	0.6389	0.6199	0.7998
48-<51	0.6736	0.6549	0.8299
51-<54	0.7066	0.6891	0.8518
54-<57	0.7318	0.7164	0.8617
57-≤60	0.7670	0.7538	0.8755
N	18384	16641	1743

Source: Derived from Nepal Fertility, Family Planning and Health Survey (NFHS) 1991.

Table 3: Coefficients for Hazard Models with the Monthly Probability of Birth as the Dependent Variable

Variables	Model1	Model2	Model3	Model4
Duration since index birth (months)				
9-<12	ref.	ref.	ref.	ref.
12-<15	0.3257*	0.4603*	0.3257*	0.4603*
15-<18	0.5665*	0.7954*	0.5567*	0.7954*
18-<21	1.0533*	1.3287*	1.0539*	1.3288*
21-<24	1.6390*	1.9209*	1.6401*	1.9211*
24-<27	1.6020*	1.9096*	1.6029*	1.9098*
27-<30	1.6789*	2.0001*	1.6794*	2.0004*
30-<33	1.8087*	2.1406*	1.8092*	2.1410*
33-<36	1.9315*	2.2585*	1.9322*	2.2590*
36-<39	1.9559*	2.2877*	1.9569*	2.2884*
39-<42	1.8269*	2.1910*	1.8286*	2.1919*
42-<45	1.8858*	2.2474*	1.8878*	2.2485*
45-<48	1.7280*	2.0688*	1.7304*	2.0702*
48-<51	1.6531*	2.0191*	1.6551*	2.0206*
51-<54	1.5741*	1.8870*	1.5759*	1.8887*
54-≤60	1.3429*	1.6968*	1.3432*	1.6978*
Characteristics of the Index Child				
Survival status				
Alive	ref.	ref.	ref.	ref.
Dead	0.8182*	2.2967*	0.7001*	2.1827*
Parity				
First birth	ref.	ref.	ref.	ref.
Low	-0.2248*	-0.2135*	-0.2248*	-0.2129*
High	-0.3586*	0.3425*	0.3552*	0.3391*
Previous interval				
Short	0.0730*	0.0739*	0.0711*	0.0727*
Medium	ref.	ref.	ref.	ref.
Long	-0.1654*	-0.1615*	-0.1658*	-0.1619*
Sex				
Male	-0.1680*	-0.1694*	-0.1910*	-0.1937*
Female	ref.	ref.	ref.	ref.
Characteristics of the Mother				
Age at index birth				
Less than 25	0.2588*	0.2602*	0.2601*	0.2617*
25 to 35	ref.	ref.	ref.	ref.
More than 35	-0.5212*	-0.5235*	-0.5212*	-0.5236*
Residence				
Urban	-0.1709*	-0.1681*	-0.1740*	-0.1705*
Rural	ref.	ref.	ref.	ref.

Variables	Model1	Model2	Model3	Model4
Education				
None	ref.	ref.	ref.	ref.
Primary	0.0809	0.0795	0.0846	0.0830
Secondary	-0.0696	-0.0755	-0.0681	-0.0740
Religion				
Hindu	ref.	ref.	ref.	ref.
Buddhist	0.0812+	0.0805+	0.0804+	0.0796+
Muslim	0.1610*	0.1683*	0.1586*	0.1666*
Others	0.3101*	0.3165*	0.3056*	0.3130*
Time-Dependent Effects				
Dead*Duration(9-<12)		ref.		ref.
Dead*(12-<15)		-0.5413*		-0.5413*
Dead*(15-<18)		-1.0233*		-1.0217*
Dead*(18-<21)		-1.3595*		-1.3555*
Dead*(21-<24)		-1.4220*		-1.4150*
Dead*(24-<27)		-1.6501*		-1.6437*
Dead*(27-<30)		-1.8098*		-1.8070*
Dead*(30-<33)		-1.9736*		-1.9724*
Dead*(33-<36)		-1.8759*		-1.8730*
Dead*(36-<39)		-1.9610*		-1.9586*
Dead*(39-<42)		-2.5467*		-2.5396*
Dead*(42-<45)		-2.4398*		-2.4326*
Dead*(45-<48)		-1.9855*		-1.9777*
Dead*(48-<51)		-2.3726*		-2.3682*
Dead*(51-<54)		-1.7181*		-1.7177*
Dead*(54-≤60)		-2.4495*		-2.4572*
Interaction Effects			0.2296*	0.2150*
Male*Dead				
Constant	-5.0620*	-5.3597*	-5.0511*	-5.3498*
- Log likelihood	31367	31243	31363	31239
Likelihood ratio	2717.9*	2662.3*	2737.9*	2664.2*
Likelihood ratio test	ref.	248.2*	8.48*	255.7*
Joint Likelihood ratio of all coefficients being zero (Model1)				
Duration since index birth	2480.44*			
Parity	64.88*			
Previous interval	41.36*			
Age at index birth	175.48*			
Education	4.5+			
Religion	18.39*			

Source: Derived from Nepal Fertility, Family Planning and Health Survey (NFHS) 1991.
*p ≤ .05, +p ≤ .10.

Figure 1: **Estimated Hazard Rates of Subsequent Birth**

(months after the birth of the index child)

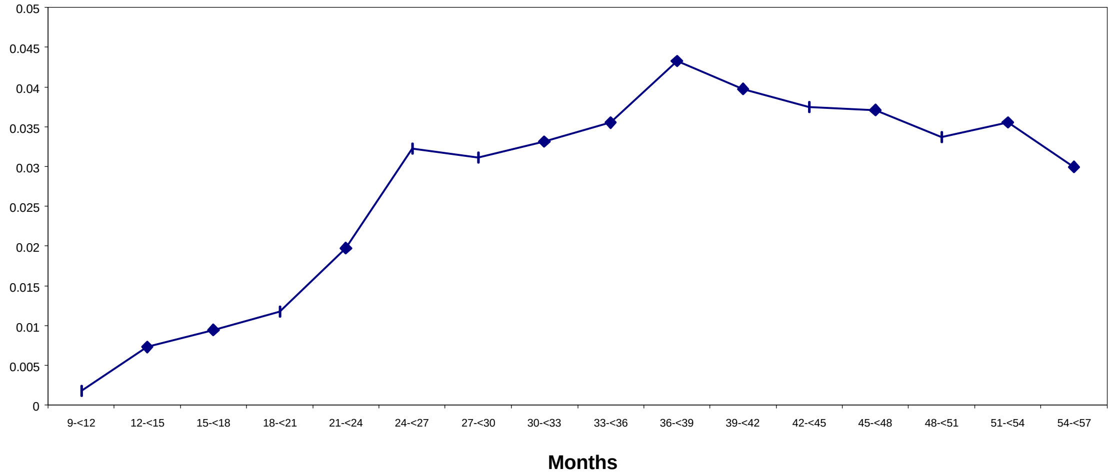

Figure 2: **Estimated Hazard Rate Ratios of Subsequent Birth between Survival Status (Dead vs Alive)**
(months after the birth of the index child)

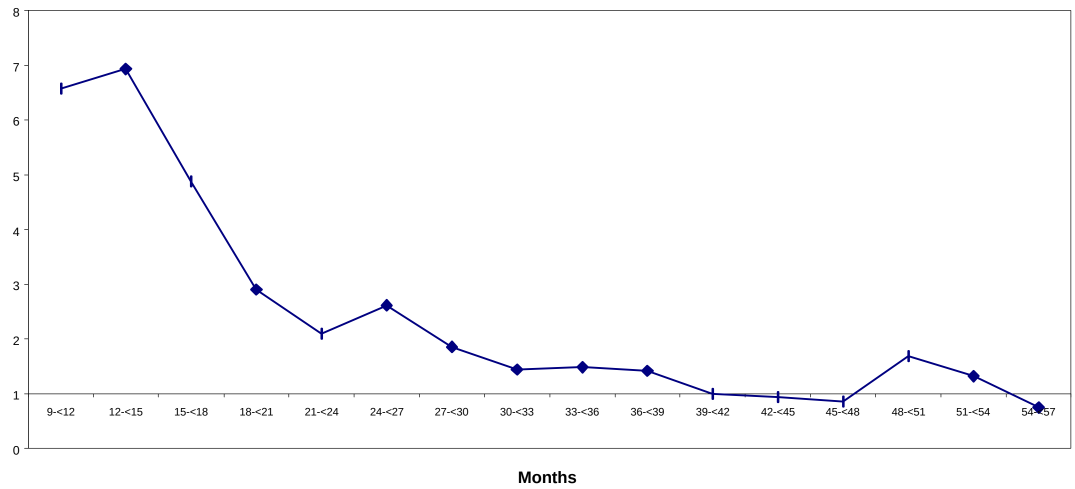

Figure 3: **Estimated Hazard Ratios of Subsequent Birth**

(months after the birth of the index child)

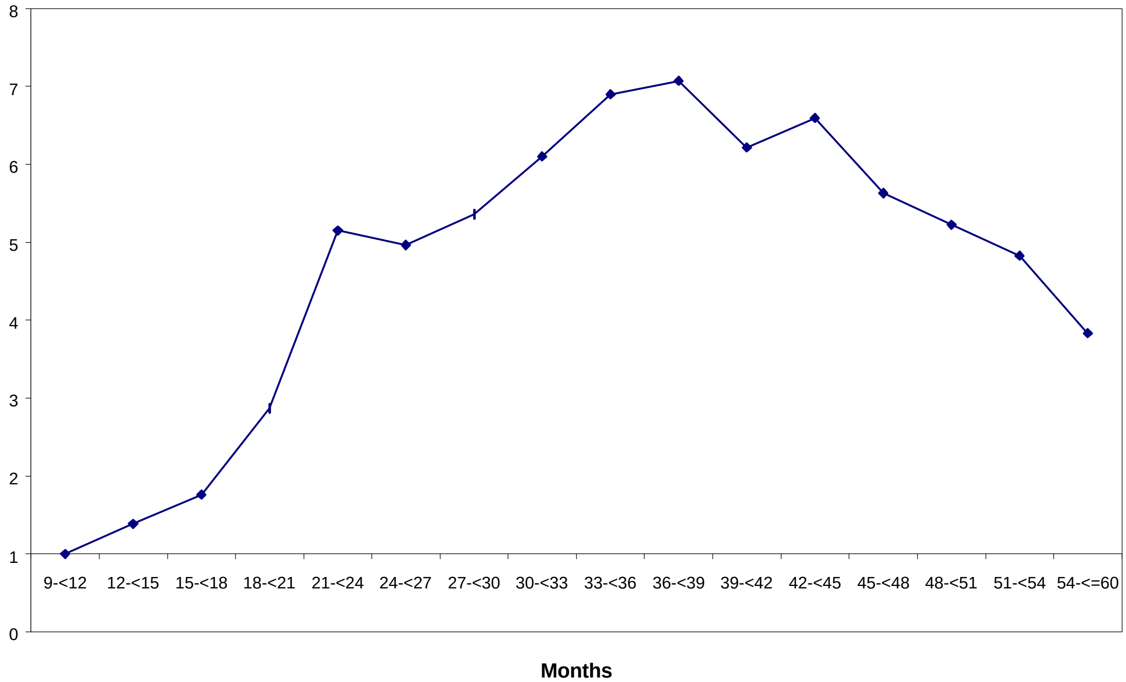

Figure 4: **Estimated Hazard Ratios of Subsequent Birth between Survival Status (Dead vs Alive)**
(months after the birth of the index child)

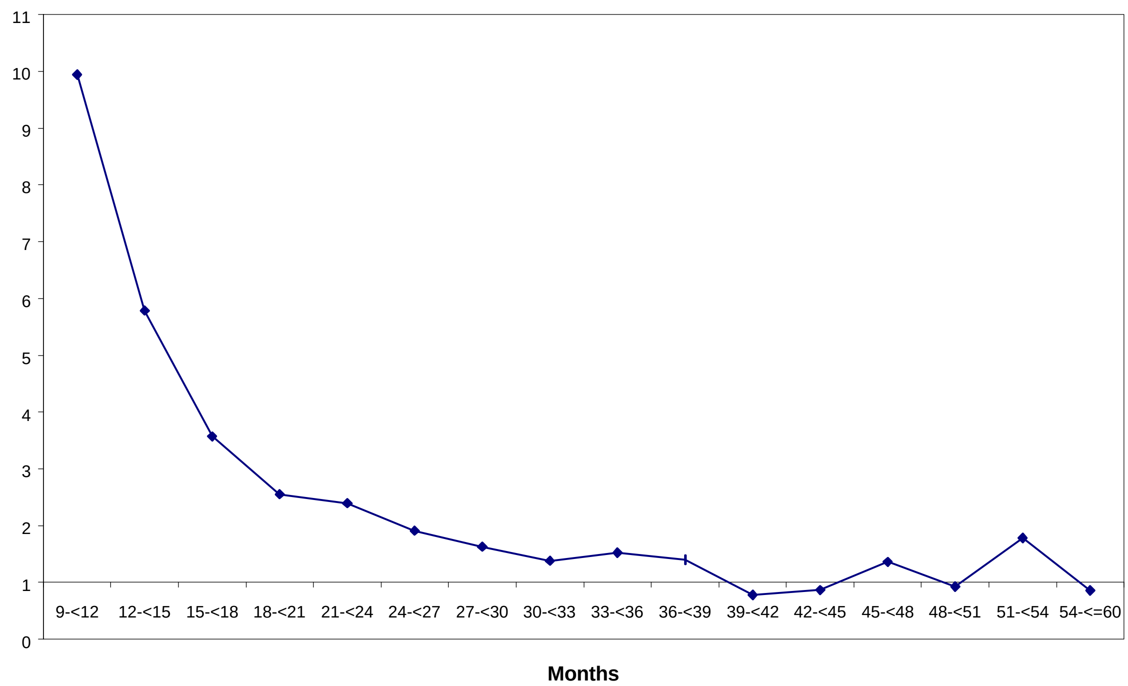

Figure 4: **Estimated Hazard Ratios of Subsequent Birth between Survival Status (Dead vs Alive)**
(months after the birth of the index child)