Effects of Birth Interval on Child Mortality: Evidence from a Sequential Analysis

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Abstract

Unlike most existing studies, this paper examines the effects of birth interval on child mortality in a sequential framework. Birth spacing is captured by the length of time since the birth of the last child and the time varying covariates identifying the arrival of a younger sibling during any month after the birth of the present child. We use an instrumental variable method to reduce the endogeneity bias and compare the hazard estimates of child survival with and without instruments for birth spacing. These instrumented sequential results not only reaffirm the static inverse relationship, but also emphasize that the inverse relationship between birth interval and child mortality crucially depends on both the gender and the birth order of the child.

1. Introduction

Quantity-quality tradeoff lies at the heart of the Beckerian models, which among other things provides a rationale for an inverse relationship between fertility and mortality. The relationship between fertility and mortality is central to many explanations of the demographic transition and important for population programmes in low-income countries, often characterized by imperfections in labour and capital markets. In the presence of these constraints, household income plays a crucial role in explaining child quality (Becker 1991) so that more siblings mean less resource per consumption unit. Thus children will do better when accompanied by siblings with fewer intrinsic advantages. Moreover, for a society with pro-male bias (Behrman et al. 1982; Sen and Sengupta 1983), children with more sisters will be better off than children with more brothers. One can also consider other possible demographic reasons affecting this negative relationship. For example, as fertility falls, pattern of family formation (e.g., relating to birth spacing,[1] parity and mother's age at birth), alters in ways beneficial to maternal and child health (Trussell 1988; Potter 1988; Hobcraft 1987).[2]

There is a long tradition of investigating the relationship between fertility and mortality in low-income countries. While Benefo and Schultz (1996) examine the effects of child mortality on fertility, LeGrand and Phillips (1996) report that higher total fertility reduces mortality in rural Bangladesh, though the effect was not very strong.[3] Some others have considered the effects of birth interval on child mortality. For example, Curtis, Diamond and McDonald (1993) report significant effects of birth interval on postneonatal mortality in Brazil. Choe, Diamond, Kim and Steele (1998) compare the effects of son preference on child mortality in Bangladesh, Egypt and South Korea and find indirect evidence that shorter birth spacing leads to higher mortality. However most existing analyses of the relationship between fertility and mortality have been done in a static framework. Also most of these studies do not address the important problem of simultaneity between fertility (or birth interval) and child mortality, which is likely to bias the available estimates.[4] The present paper attempts to address these issues while examining the effects of birth interval on child survival.

The empirical analysis of the present paper is based on the complete birth history data obtained from the 1992–93 National Family Health Survey (NFHS) from West Bengal. Since the relationship between birth interval and child mortality is in nature, we construct a sequential birth history of the sample women aged 13 to 49 years and construct birth history of children born to this women (where age is right censored at 60 months) that records the arrival of younger siblings during each month of a child's life. We estimate duration models of time to death *separately* for male and female children in terms of variables indicating birth interval, among other individual, parental/household and sibling characteristics. In particular, two types of birth spacing variables are included; namely, time since the last birth (prior birth spacing) and time varying covariate indicating arrival of younger siblings (indicators of posterior birth spacing) during each month of first sixty months of a child's life. We first estimate a conventional child survival function which depends, among other things, on these spacing variables, treating them as exogenous.[5] However, the household chooses birth spacing (and number of children) to maximize the present value of income produced by all family members. This income stream depends on the survival prospects of the children. Consequently the optimal values of birth interval (as well as number of children) will depend in part on the values of the error term in the mortality equation, giving rise to a problem of endogeneity. We use an instrumental variable method to correct for the resultant endogeneity bias of the uninstrumented estimates of mortality hazard.[6] To this end, we generate instruments for birth spacing variables, which are then included in the subsequent child survival function. We compare these estimates with and without instruments to examine the effects of simultaneity bias on child survival.

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1 In addition to reduce maternal depletion, longer birth interval also reduces the competition for limited parental care and resources in societies with credit market imperfection, thus improving the allocation of resources to the current child.

2 Bongaarts (1987), however, challenged the basis of this inverse relationship.

3 Several plausible factors were highlighted for the unexpected underestimation of the negative effects of fertility on child mortality, including the experimental design of the data from Matlab project and relatively short period of the study.

4 One important exception is Bhargava (2003), who makes a serious attempt to address the problem of simultaneity arising from the inclusion of number of older boys and older girls of the context child while analysing child survival in the Indian state of Uttar Pradesh. The analysis also includes birth interval as an explanatory variable though it was treated as a purely exogenous variable.

5 Other explanatory variables are chosen very carefully to avoid the problems of simultaneity. See further discussion in section 2.

6 We are aware of the possible problem caused by weakly correlated instruments and take care to generate the best instruments from the available information. We have also attempted to resolve this issue elsewhere by determining birth interval and child mortality as correlated processes (Makepeace and Pal, 2004).

The paper is developed as follows. Section 2 explains the data and develops the empirical child survival function while section 3 discusses the results. Section 4 concludes.

Table 1. Comparison of West Bengal with Important Indian States

States	Popln. (in mn) 1991	Female literacy Age 7+ 1991	Female labour participation 1991	Total fertility rate	Infant Mortality Rate per 1000 1990-92	Death rate, age 0-4. 1991 (per 1000) Female	Male
Kerala	29	86.2	12.8	1.8	17	4.1	4.5
Punjab	20	50.4	2.8	3.1	57	18.4	15.6
Haryana	16	40.5	6.0	4.0	71	23.8	22.3
Maharashtra	78	52.3	26.5	3.0	59	16.7	15.9
AP	67	32.7	30.1	3.0	71	20.2	22.3
Tamil Nadu	56	51.3	25.1	2.2	58	15.3	16.9
WB	68	46.6	8.0	3.2	66	20.8	20.4
India	846	39.3	16.0	3.6	80	27.5	25.6

Note: AP: Andhra Pradesh; WB: West Bengal
Source: Drèze and Sen (1995); Government of India web site: www.ni c.in/mohfw/popindi.html.

2. Data and Methodology

India is an interesting case in point. Child mortality rates in general and especially for girls are among the highest in the world.[7] There is also an interesting regional variation within the country. Table 1 compares West Bengal's demographic performance with important Indian states in 1991. Female mortality rate in the 0 to 4 years age group in 1991 was lower than the male mortality rate in the southern states of Andhra Pradesh, Kerala and Tamil Nadu, but higher in most other major states.[8] Our sample is drawn from the eastern Indian state of West Bengal. In the postindependence period, West Bengal started its economic development in a relatively good position among the Indian states, as reflected in its high rate of urbanization, strong industrial infrastructure and very high productivity of land. However, by 1967–68 the incidence of rural poverty was above-average in the state, and the situation did not improve perceptibly in the 1980s. For example, though the infant mortality rate (IMR)[9] in rural West Bengal has declined between 1981 and 1990, the state's own rate of decline in the 1980s was not much faster than the Indian average; in fact, it was surpassed or equalled by Bihar, Uttar Pradesh, Gujarat, Punjab, Kerala and Tamil Nadu (Sengupta and Gazdar 1997), thus justifying our interest to study the case of West Bengal. Even though West Bengal has moderate levels of female literacy among the Indian states, use of modern nonterminal methods of contraception remained rather low (Pal and Makepeace 2003). While about 31% of rural and 24% of urban couples were sterilized, as many as 45% of rural and 40% of urban women in NFHSS 1992–93 sample did not use any contraception. This means that only about 18.5% of rural and 24.7% of urban eligible women were currently using some modern contraception (e.g., pills, IUD, condoms, etc.).

The focus of the Indian family welfare program on sterilization as reflected in the NFHS 1992–93 has been considered to be unsatisfactory, and subsequently a new draft plan for the family welfare

7 Infant mortality rate in 1992 was 79 in India as against 18 in Sri Lanka, 31 in China, 13 in South Korea and 26 in Thailand per 1000 live births in the year.

8 Though the female mortality rates are generally lower in the Western countries.

9 Number of infants who die before reaching one year of age, expressed per 1000 live births in a given year.

programme was undertaken in 1994. The new draft focuses, among other key issues, on method-specific contraceptive targets and incentives and expanding the use of male and reversible contraceptive methods (for spacing births). We are, however, not aware of any economic analysis examining the effects of birth spacing and child survival in India, which justifies our interest in the subject.

2.1. Data

We use the National Family Health Survey (NFHS) 1992–93 household-level data[10] from rural and urban West Bengal. This allows us to construct a complete birth history for each woman aged 13-49 years. Given that the death rate in our sample tails off from age five onwards, we focus on children aged five years or below. For each child in the sample, t starts with a value of zero at birth and is right censored at 60 months. There are 12,902 children in our sample, of whom 51% are male.[11] Considering the residential location, 81% male and 82% of the female children in our sample came from rural areas of the state. About 14% of both rural male and female children died before reaching the age of 60 months, while the corresponding proportion was lower for children living in urban location (10% for female and 11% for male).

2.2. Empirical Child Survival Function

The indicator of child health in our analysis is the time to death. In general, the time to death of the i-th child, T, will depend on a vector of characteristics, X_i.

Among other things, we highlight the role of family resources. Since older first-time mothers and literate mothers tend to be more educated[12] and from higher income families, we use the age at first birth and literacy as proxies for income and wealth. We have obtained the first principal component of a number of asset variables and use this as a measure of wealth.

Castes of the Hindu households are considered to be good indicators of household wealth in the Indian context. Hence, we include controls for the lower castes.

The provision of public services like water, sanitation and health depends on the residential location. Since the availability of these services influences child survival, we include dummies for whether the household lives in a rural area and whether the household lives in a backward area. Since there may be gender differences, we include a gender dummy (in the pooled regression only) and whether the first child is female as an instrument for the presence of female children.

Variables capturing spacing between consecutive siblings may affect child health in several ways. First, shorter birth spacing would imply more maternal depletion, and therefore a limited ability of the mother to take care of the current child (including breast feeding). In addition, shorter birth spacing would imply more children, and therefore a greater strain on other parental resources per consumption units. In other words, parents can devote more time and effort to bring up a child if there is longer prior and posterior birth spacing, which would also help to limit the family size.[13] To this end, we include a variable measuring the time since the birth of a previous sibling. We also

10 The second NFHS undertaken in 199–99 is designed to strengthen the database further and facilitate implementation and monitoring of population and health programmes in the country. The principal objective of NFHS-2 remained as before, though it included some additional information relating to measurements of the nutritional status (e.g., height and weight) of all eligible women, blood test of women and children for haemoglobin. However, the information regarding birth interval and child survival that we use remained very similar.

11 This basic data-set is expanded to include monthly birth history of younger siblings per woman in our sample and the size of this expanded sample is 620469.

12 Information about the father was collected from the woman concerned. There were lots of missing as well as inconsistent values for father's age. Secondly, most fathers were literate, and hence it was causing problems of convergence. Thus, we could not include comparable characteristics of the father as we did for the mother.

13 There are other health risks of shorter birth spacing as well. For example, it may increase the likelihood of having premature babies or transmission of diseases because of crowding in the family.

include time varying variables that take the value 0 until a subsequent child is born and the value 1 thereafter.[14]

Among other sibling characteristics, the gender of the first child cannot be correlated with the innate ability of the second child, and can therefore be used as an exogenous variable in determining child survival. Similarly, without much problem, one can include if the first child died to indicate familial genetic problems, if any. The latter may take account of "death clustering" such that families experiencing child death may have shorter birth intervals (Dasgupta 1997). We have also included a variable to indicate if the current child is one of twins.[15] The latter can be treated as another health variable since it will be associated with factors such as low birth weight (although competition for limited parental resources will also play a role).

Finally, cultural preferences for sons in the Indian society are found to be important in birth spacing, and therefore in child survival. Hence one needs to examine the important role of gender in this respect. We start our analysis by considering a pooled regression with gender interaction terms and find significant differences by gender with respect to many explanatory variables. Then we proceed to consider separate male-female hazard functions for child survival.

The full list of characteristics employed in the empirical analysis and their definitions are given in Appendix 1.

The relationship between time to death and the vector of characteristics is commonly specified in terms of the hazard rate. The hazard rate, $\lambda_i(t)$, shows the rate at which individuals die at age t given that they have lived to be t years old. We use Cox's proportional hazard model[16] so that the hazard rate for t is:

$$\lambda_i(t) = \exp(X_i\beta)\lambda_0(t)$$

The hazard has "baseline" component, $\lambda_0(t)$, common to all individuals with a value that depends only on t. Systematic differences between individuals ($i \neq j$) shift the hazard. For example, when only the value of the k-th regressor is different, then

$$\lambda_i(t)/\lambda_j(t) = \exp(X_{ki} - X_{kj})\beta$$

A positive value for β implies that an increase in the value of a regressor will increase the hazard rate and, hence, lower the time to death.

3. Empirical Analysis

The main purpose of this section is to present the hazard estimates of time to death among young male and female children in West Bengal. We begin by examining the empirical distribution of time to death (section 3.1), and then consider the estimates of child survival hazard functions (section 3.2).

3.1. Kaplan-Meier Survivor Functions

Let us start with a description of the empirical distribution of the duration of child survival using Kaplan-Meier estimates of the survivor function. The survival function $S(t)$ shows the proportion

14 Both these indicators of birth spacing are clearly endogenous to child survival. We deal with this problem by producing instruments for these variables. See further discussion in section 3.

15 One, however, needs to be careful about the treatment of the twins and the corresponding birth order since birth order in our data-set is recorded in a continuous fashion, without taking account of the twin birth. Here, we have given the second born twin the same birth order as the first born.

16 Cox's proportional hazard specification is a semiparametric model of duration that is analytically tractable and places relatively few restrictions on the hazard. It also allows the use of time-varying covariates that plays an important role in our analysis.

of children that survive until period *t*. The estimates are summarized by gender in Table 2 for (a) first-born children, (b) "other" children[17] and (c) all children. To simplify the table, we present the estimates at six points corresponding to birth[18] and each year up to the age of five.

We focus on the implications of these survival rates for the corresponding hazard rates. For example, the survival rate of a first-born female child is 0.9219 at birth; the corresponding hazard rate at birth turns out to be 0.0781. For first born male child, however, the hazard rate at birth is 0.0921, higher than that of girls by about 1 and a half percentage point. Similarly, the hazard rate at birth is higher for "other" male children (0.0574 as against 0.0523), though the margin is much

Table 2. Kaplan-Meier Survivor and Hazard Functions

	First-born	Middle-order & youngest	All children
Periods (years)	**Male children**		
0	0.9079 (0.0921)	0.9426 (0.0574)	0.9322 (0.0678)
1	0.8631 (0.05)	0.9037 (0.04)	0.8919 (0.0432)
2	0.8532 (0.0115)	0.8926 (0.0123)	0.8810 (0.0122)
3	0.8440 (0.0108)	0.8852 (0.0083)	0.8732 (0.0089)
4	0.8403 (0.0044)	0.8764 (0.01)	0.8659 (0.0084)
5	0.8383 (0.0024)	0.8729 (0.0039)	0.8629 (0.0035)
	Female children		
0	0.9219 (0.0781)	0.9477 (0.0523)	0.9400 (0.06)
1	0.8813 (0.044)	0.9049 (0.045)	0.8979 (0.0448)
2	0.8714 (0.0114)	0.8932 (0.013)	0.8868 (0.0124)
3	0.8629 (0.0098)	0.8794 (0.0155)	0.8746 (0.0138)
4	0.8571 (0.0068)	0.8716 (0.0089)	0.8673 (0.0083)
5	0.8551 (0.0023)	0.8662 (0.0062)	0.8629 (0.0051)
Logrank test			
Chi-square (1)	2.25	0.53	0.05
P-value	0.1335	0.4651	0.8171
Wilcoxon test			
Chi-square (1)	2.53	0.26	0.24
P-value	0.1120	0.6097	0.6242

Note: Numbers in the parentheses show the corresponding hazard rates.

smaller; we also note that compared to first-born children, the hazard at birth is generally lower for the nonfirst-born children of a given gender. Thereafter, the hazard rate decreases with the passage of time for both male and female children (irrespective of their birth orders) such that the male-female differential also vanishes. For example, at t=1, the male-female differential is just about one percentage point for first-born, while at t=4, female hazard is slightly higher for the first-born child. This pattern can be explained by various exogenous factors, which takes us to the determination of hazard rates in terms of a multivariate regression framework, to which we now turn.

We also test for the equality of survivor functions between male and female children of different birth orders. Two standard rank tests – namely, the logrank and Wilcoxon[19]– are applied (see Table

17 It is important to distinguish first-born children from other nonfirst-born ones since mortality risks are found to be different for these two groups of children (e.g., see Muhuri and Preston 1991).

18 When we estimate the statistical models, we assume that people who die at birth are one day old when they die.

19 Logrank test is appropriate when the hazard functions are thought to be proportional across gender, while Wilcoxon test is appropriate when the hazard functions vary nonproportionally.

Table 3: Sample Characteristics – Means and Standard Deviations (in the parentheses) for the Time-Varying Data-set

Variables	Male	Female	All
AGEMUM1	17.57 (2.93)	17.63 (2.96)	17.6 (2.95)
LITMUM	0.38 (0.48)	0.38 (0.48)	0.38 (0.48)
TWIN	0.01 (0.09)	0.008 (0.09)	0.01 (0.09)
FSTFEMALE	0.36 (0.48)	0.64 (0.48)	0.5 (0.5)
FIRSTDIE	0.18 (0.39)	0.18 (0.38)	0.18 (0.38)
PREVHAT	38.67 (8.07)	37.65 (7.85)	38.18 (8.05)
YHAT1	0.26 (0.44)	0.28 (0.45)	0.27 (0.44)
BORDER	2.87 (1.86)	2.88 (1.86)	2.87 (1.86)
AGLAND	0.53 (0.50)	0.53 (0.49)	0.53 (0.50)
PUCCA	0.16 (0.36)	0.15 (0.36)	0.15 (0.36)
RADIO	0.41 (0.49)	0.41 (0.49)	0.41 (0.49)
TELE	0.13 (0.34)	0.13 (0.34)	0.13 (0.34)
PCASSET	0.04 (1.01)	0.02 (1.02)	0.03 (1.02)
SC	0.10 (0.29)	0.10 (0.30)	0.10 (0.29)
ST	0.05 (0.22)	0.06 (0.23)	0.05 (0.22)
MUSLIM	0.31 (0.46)	0.29 (0.46)	0.30 (0.46)
RURAL	0.80 (0.40)	0.81 (0.39)	0.81 (0.39)
BACKWD	0.25 (0.43)	0.25 (0.43)	0.25 (0.43)
Observations	315784	304685	620469

2) by gender for first-born, nonfirst-born (middle-order and youngest children) and also for all children taken together. We cannot reject the hypothesis that the survivor functions are the same for the nonfirst-born and all children; according to the Wilcoxon test, we can, however, reject the null hypothesis of equality of survivor functions at about 11% level for the first-born children in our sample.

3.2. Determinants of Child Survival

There are two methodological issues to be addressed here. The first one relates to the gender difference in child survival, if any. We start our analysis by considering the pooled sample of male and female children to examine how the exogenous variables may explain the gender differences in child survival for first-born and other (middle order and youngest) children.[20] In addition to variables reflecting family endowment, household assets and sibling characteristics (as described above), we include a set of gender interaction terms. We construct a likelihood ratio statistic to test the joint significance of the gender interaction terms. Both LR statistics are statistically significant,[21] suggesting that there is significant difference by gender with respect to various exogenous explanatory variables in our sample. Hence we focus on the estimates obtained from separate male-female hazard functions for child survival.

The second issue is more complicated and relates to the possible simultaneity of the variables related to birth spacing in determining child survival. First we estimate a survival hazard model where birth spacing is treated as a purely endogenous variable. We then adopt an instrumental variable approach where we first generate appropriate instruments of these indicators of birth spacing and

20 There are some differences in the specifications for the regressions for the first-born children and other children. For obvious reasons, birth order, characteristics of the first-born sibling and the prior birth spacing instrument are not included in the equation for the first-born children.

21 The significance of gender differences are also confirmed when a Weibull model is used in place of the Cox model.

then include these instruments, among others, to determine the child survival hazard function. Finally we compare these two sets of estimates, with and without instruments, to find out the effects of simultaneity bias on child survival.

Table 4. Uninstrumented Hazard Estimates for Child Survival
(Where time since last birth is PREVIOUS and presence of younger siblings is YOUNG1)

	Male		Female	
Variables	**First born**	**Middle-order and youngest**	**First born**	**Middle-order and youngest**
AGEMUM1	-0.04 (1.85)*	0.003 (0.22)	-0.13 (5.05)**	0.01 (0.38)
LITMUM	-0.26 (1.85)*	-0.33 (3.02)**	-0.25 (1.68)*	-0.41 (3.69)**
TWIN	1.53 (3.71)	1.35 (5.16)**	1.76 (4.42)**	0.87 (3.43)**
FSTFEMALE	-	-0.01 (1.20)	-	0.01 (0.08)
FIRSTDIE	-	0.17 (1.77)*	-	0.31 (3.37)**
PREV1	-	-0.03 (9.35)**	-	-0.03 (8.08)**
BORDER	-	-0.04 (1.42)	-	-0.02 (0.60)
PCASSET	-0.15 (2.05)*	-0.12 (2.27)*	-0.14 (1.81)*	-0.09 (1.55)
SC	0.02 (0.09)	0.4 (3.13)**	-0.13 (0.65)	0.19 (1.49)
ST	0.4 (1.82)*	0.39 (2.36)*	-0.39 (1.36)	-0.03 (0.17)
MUSLIM	0.13 (0.89)	-0.19 (1.76)*	-0.03 (0.19)	0.01 (0.13)
RURAL	0.02 (0.11)	-0.04 (0.33)	0.02 (0.09)	0.20 (1.54)
BACKWD	-0.17 (1.15)	0.29 (2.94)**	0.32 (2.21)*	0.02 (0.20)
YOUNG1	0.28 (0.81)	0.35 (1.72)*	0.41 (1.21)	0.69 (3.71)**
Log-likelihood	-2279.5783	-4572.1572	-1886.8024	-4592.0449
Chi-square	76.77	236.53**	130.21**	208.15
Number of *subjects* [1]	1955	4581	1856	4416

Note: [1] Note that the total number of observations in each case was at most 60 times the number of subjects in each category. Numbers in the parentheses denote the corresponding t statistics. '*' denotes that the variable is significant at10% or lower level , while '**' denotes the same at 1% level.

Table 5. Hazard Estimates of Birth Spacing for All Children

(1) Variables	(2) Mean (standard deviation)	(3) Coefficient (T-ratio)
INTERCEPT	-	-3.44 (41.991)**
AGEMUM1	17.64 (2.98)	-0.04 (10.453)**
PUCCA	0.15 (0.35)	-0.23 (6.155)**
TELE	0.12 (0.33)	-0.25 (5.666)**
RADIO	0.4 (0.49)	-0.11 (4.648)**
SC	0.10 (0.30)	-0.08 (2.164)*
MUSLIM	0.31 (0.46)	0.14 (5.716)**
RURAL	0.82 (0.39)	0.14 (4.462)**
FIRSTDIE	0.22 (0.41)	0.13 (5.210)**
MALE	0.51 (0.50)	-0.04 (2.003)*
Log-likelihood	-	-153076.977
Chi-square	-	746.55**
In a	-	0.096 (12.590)**
Number of observations	12902	12902

Note: These estimates are obtained by starting with the full model and then excluding insignificant regressors (namely, LITMUM, AGLAND, ST, BACKWD, FIRSTFEM) at successive steps until all remaining regressors are statistically significant. Number in the parentheses denote the corresponding t statistics. '*' denotes that the variable is significant at 10% level or lower and '**' denotes the same at 1% level or lower.

Table 6. Hazard Estimates for Child Survival (with Instruments)

	Male		Female	
Variables	**First born**	**Middle-order and youngest**	**First born**	**Middle-order and youngest**
AGEMUM1	-0.03 (1.672)*	0.01 (0.553)	-0.13 (5.373)**	0.03 (1.203)
LITMUM	-0.26 (1.830)*	-0.3 (2.796)**	-0.24 (1.675)*	-0.43 (3.879)**
TWIN	-1.01 (1.573)	0.9 (2.628)*	0.11 (0.187)	1.04 (3.259)**
FSTFEMALE	-	-0.1 (1.117)	-	0.03 (0.354)
FIRSTDIE	-	0.17 (1.687)*	-	0.27 (2.609)**
PREVHAT	-	-0.01 (0.987)	-	-0.03 (1.705)*
BORDER	-	-0.03 (1.237)	-	-0.01 (0.498)
PCASSET	-0.14 (1.926)*	-0.07 (0.916)	-0.13 (1.656)*	0.01 (0.150)
SC	0.007 (0.035)	0.36 (2.775)**	-0.14 (0.692)	0.17 (1.317)
ST	0.4 (1.823)*	0.39 (2.398)*	-0.39 (1.334)	0.004 (0.023)
MUSLIM	0.1 (0.686)	-0.18 (1.526)	-0.06 (0.396)	-0.02 (0.211)
RURAL	0.01 (0.047)	-0.12 (0.853)	0.001 (0.007)	0.02 (0.118)
BACKWD	-0.17 (1.188)	0.28 (2.796)**	0.31 (2.173)*	-0.001 (0.015)
YHAT1	2.85 (4.703)*	0.62 (2.058)**	2.08 (3.779)*	0.52 (1.994)*
Log-likelihood	-2271.615	-4629.2518	-1881.0489	-4639.0224
Chi-square	92.69**	122.34**	141.72**	114.09**
Number of *subjects* [1]	1955	4581	1856	4416

Note: [1] Note that the total number of observations in each case was shown in Table 3. Number in the parentheses denote the corresponding t statistics. '*' denotes that the variable is significant at10% or lower level while '**' denotes the same at 1% level. Instruments for time since last birth is PREVHAT and that for arrival of younger siblings is (YHAT1).

3.2.1. Hazard Estimates without Instruments

These traditional estimates are summarized in Table 4. A positive coefficient suggests that the hazard rate of dying increases with the increase in the corresponding variable while a negative coefficient implies the opposite.

The hazard of having a subsequent child is lower for older first time mothers and also literate mothers. However, being one of the twins and shorter prior spacing enhance the hazard of mortality. Arrival of the first younger sibling is only significant for the higher order children, but not for the first-born child. Gender of the first child is not significant for any child, while death of the first child significantly enhances the mortality risks of the higher order children. Household wealth significantly lowers the hazard, but primarily among young boys in our sample.

We next move on to produce the instrumented estimates of child survival with a view to compare them with the uninstrumented estimates.

3.2.2. Derivation of Instruments

In household models, parental optimisation not only determines quantity and quality of children, but also the spacing between children. Thus the fundamental problem is to find suitable instruments that reflect the spacing with the previous and subsequent siblings. In order to redress this important endogeneity problem, we take a two-step approach.

We first estimate a Weibull survival model with birth spacing (estimated as time to the birth of the next child) as the dependent variable using the largest possible sample of children. We then estimate the typical interval between births as the median of the estimated distribution. We chose Weibull model because it is relatively easy to produce estimates of the median.

The set of explanatory variables includes parental and household characteristics (AGEMUM1, LITMUM, SC, ST, MUSLIM), household assets (AGLAND, PUCCA, RADIO, TELE), characteristics of the children already born (FIRSTFEM, FIRSTDIE), and the gender of the current child (MALE).[22] The choice and use of contraceptives are important determinants of birth spacing in many cases, though it is a choice variable for the couple in question and therefore, raises the question of endogeneity. Hence we use instruments that can reflect use of contraception in our sample. The religion variable, MUSLIM, is a good instrument, in that the use of modern contraception is rather limited among the Muslim couples in our sample. Mother's age is a good measure of fecundity, while mother's literacy is widely found to reduce fertility. Household assets as well as caste variables (SC, ST) instrument household wealth position in the Indian context. We also argue in our analytical model that the gender composition of the existing children (e.g., FIRSTFEM and MALE) could significantly affect parental birth spacing decisions. In this respect, we use the nontime varying sample; summary statistics are shown in column 2 of 4.

The final model specification is chosen by dropping the insignificant variables (e.g., LITMUM, AGLAND, ST, FIRSTFEM, BACKWD) in successive rounds until all the existing explanatory variables are significant at 10% or lower level. These estimates are shown in column 3 of Table 5. Mother's age and most assets variables lower the hazard of subsequent child birth. However, the hazard is higher if the couple is Muslim and the first child of the couple dies. More interestingly, we find that the hazard of subsequent child birth is lower if the current child is male, which in turn reflects some indirect evidence of son preference in birth spacing. These Weibull estimates are then used to predict the indicators of prior and posterior birth spacing to be used in the second stage child survival hazards.

Our model includes the time since the birth of a previous sibling and time varying covariates that take the value 0 until a subsequent child is born and the value 1 thereafter. The first variable is instrumented by the estimated median of time to the birth of the next sibling defined as:

$$\text{PREVHAT}=\mathit{median}=-\frac{1}{\hat{\theta}}\left(1\text{n}\left(\frac{1}{2}\right)\right)\frac{1}{\hat{a}}$$

where the Weibull hazard is $\lambda(t) = a\theta(X)^a t^{a-1}$and θ and $\hat{a}$ are the estimates of θ and a.[23]

The dates when subsequent children are born are estimated as PREVHAT in the future. The time varying covariate based on the time to the birth of the next child takes the value 0 for the period of time that is less Prevhat in the future and 1 thereafter.

We note that Prevhat is a highly nonlinear function of the regressors in the Weibull model and that the time varying covariates are not linear functions of either Prevhat or the regressors. This will ensure the identification of the coefficients in the Cox child survival model. The variable for Muslim provides a more traditional approach to identification. It is always a highly significant regressor in the birth spacing model but has very high p-values in the child survival model.

3.2.3. Hazard Estimates with Instruments

Estimates of the Cox proportional hazard model with time varying covariates are shown in Table 6 for male and female children separately.[24] The estimated coefficients of a hazard function do not give any indication of the marginal effects of the explanatory variables. Hence, we have also reported the hazard ratios in Table 7. In response to a change equal to one unit in the value of an explanatory variable, the hazard ratio for the variable concerned measures by how much the hazard rate changes in relation to the hazard rate for the original value of the variable.

22 See Appendix for the definition of these variables.

23 We chose the median rather than the mean because skewedness of the Weibull distribution.

24 We have also performed similar regressions with the Weibull specification of the hazard and found that the results are very similar to the Cox estimates. These results can be supplied at request.

Table 7. Hazard Ratios for Male and Female Children

	Male		Female	
Variables	**First born**	**Middle-order and youngest**	**First born**	**Middle-order and youngest**
AGEMUM1	0.97 (-1.612)	1.01 (0.553)	0.88 (-5.373)**	1.03 (1.203)
LITMUM	0.77 (-1.830)*	0.74 (2.796)**	0.78 (-1.675)*	0.65 (-3.879)**
TWIN	0.37 (-1.573)	2.44 (2.628)**	1.12 (0.187)	2.83 (3.259)**
FSTFEMALE	-	0.91 (-1.117)	-	1.03 (0.354)
FIRSTDIE	-	1.19 (1.687)*	-	1.31 (2.609)**
PREVHAT	-	0.99 (-0.987)	-	0.97 (-1.705)*
BORDER	-	0.97 (-1.237)	-	0.99 (-0.498)
PCASSET	0.87 (-1.926)*	0.93 (-0.916)	0.88 (-1.656)*	1.01 (0.150)
SC	1.01 (0.035)	1.43 (2.775)**	0.87 (-0.692)	1.19 (1.317)
ST	1.5 (1.823)*	1.48 (2.398)**	0.68 (-1.334)	1.004 (0.023)
MUSLIM	1.1 (0.686)	0.83 (-1.526)	0.94 (-0.396)	0.98 (-0.211)
RURAL	1.01 (0.047)	0.89 (-0.853)	0.99 (-0.007)	1.02 (0.118)
BACKWD	0.84 (-1.188)	1.32 (2.796)**	1.37 (2.173)*	0.99 (-0.010)
YHAT1	17.2 (4.703)**	1.86 (2.058)*	7.9 (3.779)**	1.69 (1.944)*
Number of *subjects* [1]	1955	4581	1856	4416

Note: [1] Note that the total number of observations in each case was sixty times the number of subjects in each category. Number in the parentheses denote the corresponding t statistics. '*' denotes that the variable is significant at10% or lower level while '**' denotes the same at 1% level.

First-born male and female children: The hazard rate of dying significantly decreases with increase in mother's age at the birth of the child and mother's literacy for both male and female children. Also, the household composite current asset variable, PCASSET, has a significant negative effect on the hazard rate for both male and female children; thus children from better off households have lower hazard in survival. First-born male children from scheduled tribe households have higher hazard rate, while the caste/religion variables are not significant for the female children. Instead, the hazard rates are significantly higher for first-born female children from backward areas of the state, while this variable is insignificant for boys.

Among various sibling characteristics, twin birth does not seem to affect the hazard rate, while the arrival of a younger sibling does have a significant impact on the hazard rate of the first-born male and female children. In particular, results suggest that arrival of younger siblings increases the hazard rate for both male and female children in our sample.

Estimates for first-born female children are generally similar to those of first-born male children. Certain differences are, however, noteworthy, especially with respect to household caste and residential location. In particular, boys from scheduled tribe households and girls from backward regions of the state have significantly higher hazard rates (and therefore significantly lower survival rates), even after controlling for household assets in our sample.

Middle order and youngest male and female children: Among parental characteristics, mother's literacy is found to play a significant role for both boys and girls, children with literate mothers having significantly lower hazard rate than those with illiterate mother. However, mother's age or composite household assets variable is not significant for male or female samples. Female children from scheduled caste families and male children from scheduled caste and scheduled tribe families have significantly higher hazard rates than the reference group of upper caste Hindus. In addition, male children from backward areas have a significantly higher hazard, though the variable is insignificant for female children.

There are also significant sibling composition effects for these groups of male and female children. Longer prior birth spacing lowers hazard rate for female children, but it is not significant for boys. However, death of the first sibling (FIRSTDIE), a twin birth (TWIN) and arrival of a subsequent sibling (YHAT1) enhance the hazard rate of both boys and girls. To conclude, besides caste and locational factors, gender difference in survival for nonfirst-born children also depends on prior birth spacing.

3.3. A Comparison of Hazard Estimates with and without Instruments

We conclude our analysis by comparing the hazard estimates with (Table 6) and without (Table 4) instruments (PREVHAT AND YHAT) of indicators of birth spacing (PREV1 AND YOUNG1). One can highlight the interesting differences in estimates with respect to these birth spacing variables. (a) For example, for first-born male children arrival of younger siblings (YOUNG1) is insignificant in the hazard regression (Table 4), while the corresponding instrumental variable YHAT1 becomes significant (Table 6). (b) For higher order children, however, prior birth interval (PREV1) is significant for both male and female children and the corresponding t-statistics seem to be spuriously high. If, however, we consider the corresponding instrumented estimates, PREVHAT is insignificant among higher order male children, while it is significant among female children (t-statistic is more reasonable too). Estimates with respect to the arrival of younger children (YOUNG1 and YHAT1) remain very similar in these two sets of estimates. Thus the relationship between birth interval and child mortality in our sequential framework not only varies between boys and girls, but also among children of a given gender, but of different birth orders.

Signs of all other regressors that are significant turn out to be similar, though t-statistics tend to be generally higher for the uninstrumented estimates. However, the values of the log likelihood function and the chi-square LR statistics are higher for the instrumented estimates, suggesting an overall improvement in the goodness of fit for the two-step instrumented estimates child survival. Taken together, there is indirect evidence that the hazard of having subsequent sibling is higher for girls (see Table 5), which in turn suggests a higher mortality risks because of shorter birth spacing following the birth of a female child. These instrumented mortality hazard estimates also suggest that shorter prior birth spacing lowers the hazard of higher order female children (the variable is insignificant for male children), while the arrival of next younger sibling enhances the hazard among all children irrespective of their gender in our sample.

4. Concluding Comments

The present paper examines the effects of birth interval on child mortality in the eastern Indian state of West Bengal. The empirical analysis is based on the sequential birth history of women aged 13-49 years. Birth spacing is captured by the length of time since the birth of the last child and the time varying covariates identifying the arrival of younger sibling after the birth of the present child. We examine the effects of birth spacing on the mortality hazard of young male and female children of different birth orders in West Bengal. In doing so, we also attempt to reduce the endogeneity bias by including instruments for birth spacing in the mortality equation and then comparing the estimates with and without instruments. It is argued here that the instrumented estimates reduce the estimation bias. These estimates suggest that longer prior birth spacing significantly lowers the mortality hazard among higher order female children. First, as the birth spacing increases, the chances of survival improve for the later-born child. perhaps because parents are able to devote more time and effort to bringing that child through his or her critical early years. However, this risk is insignificant among young boys, which perhaps indicates the importance of son preference among parents in this part of the world. Importance of son preference is also indicated by the fact that generally female children face a higher hazard of having a younger sibling than their male counterparts. The birth of the immediately next child[25] decreases the chances of survival among male and female children of

25 The effects of other subsequent children may not be pronounced because the numbers in the sample are relatively small.

any birth order, in this case because the time, resource and effort spent on the context child decline. This could also be a result of a greater maternal depletion due to shorter the spacing.

These instrumented sequential results not only reaffirm the static inverse relationship, but also emphasize that the inverse relationship between birth interval and child mortality crucially depends on both the gender and the birth order of the child. These results also question the rationale for India government's heavy-handed policy of sterilization. Instead our results emphasize the potential success of modern family planning programmes, aiming at spacing births through popularizing various nonterminal modern methods of contraception.

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Appendix – Variable Definitions

The analysis is based on the National Family Health Survey (NFHS) 1992–93 household-level data for West Bengal.

Regression variables

AGEMUM1:	Age of the mother at the birth of the first child
LITMUM:	1 if the mother is literate and 0 otherwise
BORDER:	Birth order
TWIN:	1 if the child is a twin or a triplet and 0 otherwise
FIRSTFEM:	1 if the first sibling in the family is a female and 0 otherwise
FIRSTDIE:	1 if the first sibling in the family died and 0 otherwise
PREV1:	Length of time since the birth of the immediately previous child
PREVHAT:	An instrument for PREVIOUS
YOUNG1	0 for periods before the birth of the first child after the current child 1 thereafter
YHAT1:	Instrument for YOUNG1
AGLAND:	1 if the household owns agricultural land and 0 otherwise
PUCCA:	1 if the household lives in a brick house and 0 otherwise
RADIO:	1 if the household owns a radio and 0 otherwise
TELE:	1 if the household owns a television and 0 otherwise
PCASSET:	A composite measure of household assets (the first principal component of AGLAND, PUCCA, RADIO and TELE
MUSLIM	1 if the family is Muslim and and 0 otherwise
SC	1 if the family is from a lower caste (Hindu only) and 0 otherwise
ST	1 if the family is from a scheduled tribe and 0 otherwise
RURAL:	1 if the child lives in rural areas and 0 otherwise
BACKWD:	1 if the child lives in a backward area and 0 otherwise
MALE:	1 if the child is male and 0 otherwise