Understanding the Effects of Siblings on Child Mortality: Evidence from India

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Abstract: This paper examines the efffect of siblings on child mortality in the Indian state of West Bengal arguing that prior and posterior spacing between consecutive siblings are important measures of the intensity of competition among siblings for limited resources. Parental decisions regarding spacing is endogenous to allocation of resources though available estimates of child mortality largely ignore it. To correct for this possible endogeneity bias, we allow for family specific unobserved heterogeneity and model birth spacing and child mortality as correlated processes within a sequential framework. These corrected estimates suggest: (a) the hazard of prior spacing may increase while that of posterior spacing decrease with mother's literacy and household assets. (b) the chances of child survival increase with an increase in both prior and posterior birth interval but decrease with the birth of a twin. (c) prior and posterior birth intervals have different effects on young boys and girls, which, in turn, reflect the nature of decisions made by resource constrained parents characterised by pro-male bias.

JEL Classification : D13, I12, O15

Key Words: Sibling competition, Age and gender composition, Birth spacing, Child mortality, Pro-male bias, Unobserved heterogeneity.

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1. INTRODUCTION

Children in low-income countries face much higher risks of mortality compared to their counterparts in more affluent societies. While the infant mortality rate in 1992 was 79 per thousand in India, it was only 26 in Thailand and 13 in South Korea. This disadvantage often arises from the lack of parental resources in societies characterised by credit market imperfections. The problem is further aggravated for larger families with more children as these families need to allocate limited available resources across more consumers. Even in the absence of any strategic behaviour by family members, children compete for limited parental care and resources – a notion commonly labelled as 'sibling rivalry' in economics (Garg and Morduch, 1998). The essential implication of this sibling rivalry is that sibling composition becomes crucial for determining child survival. Garg and Morduch (1998) quantified sibling composition by including number of brothers and sisters or number of older brothers and sisters into a child health function. These indicators of sibling rivalry cannot capture the age differences between consecutive siblings and thus the intensity of competition between prior and posterior siblings. By the time a child is born, an older sibling may not require extensive care from the parents and may even help parents by looking after younger siblings or supplementing family earnings. Thus another important aspect of sibling rivalry that has not been adequately taken account of is the spacing between consecutive children. Following Rosenzweig 91986), we consider a sequential framework and examine among other things role of age (e.g., prior and posterior birth spacing) and gender (e.g., gender of the first child) composition of siblings on child mortality .

There has been a long tradition of investigating the relationship between fertility and mortality in low income countries. While most researchers observe strong negative effects of fertility on child mortality (e.g., see Benefo and Schultz, 1996), LeGrand and Phillips (1996) report that the expected effect of higher total fertility on mortality reduction in rural Bangladesh has not been very strong.[1] Others have considered the effects of birth interval on child mortality. For example, Curtis, Diamond and McDonald (1993) report that shorter birth interval significantly increases post-neonatal mortality in Brazil. Choe, Diamond, Kim and Steele (1998) further 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.

Thus there is a considerable amount of demographic literature that suggests that longer birth spacing (and therefore lower fertility) is associated with lower child mortality. The explanations for this inverse relationship include, among others, less maternal depletion and more resources and parental care per child. We, in addition, argue that birth spacing reflects age difference between consecutive siblings and this therefore would enable us to capture the intensity of competition among siblings depending on their age-related needs. For example, if the spacing is less than a year, older sibling is yet to be independent when the younger sibling is born and thus both would require immediate attention of the parents, which in turn may cause a strain on parental resources. In contrast, if the spacing is three years or more, older sibling would need much less attention from the parents so that the parents can devote more of their time and energy towards the younger sibling. The relationship may however be more complex than it first appears to be. For example, longer birth spacing beyond a certain limit may give rise to some kind of intra-family conflicts including diverse child investments that do not require complementarities and maternal depletion at the other end. Some psychological literature too suggests that older child may resent the attention paid to the younger sibling if the intervals are long as the older child has had more time alone with the parents. The relationship may thus be non-linear so that longer spacing may be beneficial up to a certain extent but beyond that the relationship may be reversed.

Clearly the issue of sibling rivalry is closely related to available family resources. The Beckerian model (1991) explains the nature of parental investment in children and the quantity-quality trade-off essentially within a static framework when there are imperfections in labour and credit markets. In the presence of these constraints, children will do better when accompanied by siblings with fewer intrinsic advantages. Thus for a society with a pro-male bias (Behrman et al., 1982; Sen and Sengupta, 1983), younger children with more older sisters will be better off than those with more older brothers. Following Rosenzweig (1986), the present paper however considers a sequential framework and empirically examines the role of age and gender composition of siblings on child survival. Among other things, we

[1] 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.

consider prior and posterior spacing as indicators of age composition of siblings while gender of the first child as that of gender composition.

Our analysis is based on complete birth history data obtained from the 1992-93 National Family Health Survey (NFHS) from the eastern Indian state of West Bengal. We consider the birth history of women aged 13 to 49 years[2] belonging to households often characterised by resource constraint and son preference (Pal, 1999). Our central result is that child mortality falls when both prior and posterior spacing increases (although there is some asymmetry between these two spacing decision for each context child). Demonstrating this simple proposition however raises a number of estimation problems that we address here.

Clearly spacing decisions are endogeneous to parental allocation decisions regarding child health (that affects mortality) in a static framework.[3] In a sequential framework however there is a jointness of spacing decisions; thus an increase in parental age at the birth of the context child affects its older sibling's post-birth interval and its yonger siblings's prior birth interval. Thus even if there is no difference in child specific endowments, it is unlikely that parents could simultaneously equalise quality outcomes of all its children, even if parents care about child quality and/or are averse to inequality among sibling quality. .

We adopt a novel technique to estimate the health quality production function in a sequential framework. First given the simultaneity between sibling age differences (measured by prior and posterior interval of each child) and child health outcomes (measured by probability of a child dying or mortality hazard) in a sequential framework, we jointly determine child mortality (probit/hazard) with prior and posterior spacing decisions, allowing for different mother/family specific unobserved heterogeneity in spacing (prior and posterior) and mortality equations. We assume that the unobserved factors which partly determine birth spacing and child mortality are correlated because the same couple makes both these decisions; in other words. For example, at the same level of education and wealth and other observables, parents who choose to have shorter birth spacing intervals may also have higher death rates for their child because of this common unobserved parental effect. In this case,

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[2] Although there are younger women in our sample who have not completed fertility, our estimates seem to be robust. Not only we include mother's age at birth as a control variable in the spacing equation, but also our analysis focuses on middle-order children born to these women. In particular, we use hazard equations to determine prior and posterior spacing; oldest children are censored in the prior spacing equation while youngest ones are censored in the posterior spacing equation.

[3] Most existing studies tend to ignore this simultaneity. One important exception is Bhargava (2003) who analysed child survival in the Indian state of Uttar Pradesh. This study makes a serious attempt to address the problem of simultaneity arising from the inclusion of older boys and girls of the context child . The estimation also included birth interval as an additional explanatory variable though it was treated as a purely exogenous variable.

low values for birth spacing would be associated with high unobserved values for the propensity to die creating a correlation between birth spacing and the unobservable error term in the mortality equation. By modelling this aspect of the data generation as a common fixed effect,[4] we are able to remove the implicit bias resulting from the correlation. We also propose a recursive structure for our model, ensuring identification in the presence of the common fixed factor (See Chamberlain and Griliches, 1975). Each model (comprising of 3 correlated equations pertaining to posterior and prior spacing and child mortality) is estimated separately for male and female children in order to reduce biases due to resource allocation in favour of males.

The paper now considers the hypothesis in greater detail and describes the data and the statistical model (see section 2). The subsequent section presents and analyses the results. Findings of the paper are summarised in the final section.

2. HYPOTHESES, DATA AND METHODOLOGY

Families maximise the total income of the parents and potential children. The income of each child depends on their health which depends *inter alia* on the number of other children in the family. There are clear incentives to raise future income by having more children (which means shorter birth spacing) but the earning power of children depends on their quality, measured here by their health. The family's resources are constrained so an increase in the number of children will reduce the health of the children and their future earning capacities. This trade-off between quantity and quality (measured here by health) lies at the heart of Beckerian models and justifies our interest in testing the empirical validity of the relationship. The Beckerian school of thought is essentially static in nature and does not take account of the sequential nature of childbirth. One important extensions in this respect is provided by Rosenzweig (1986)[5] that provides the rationale for our work in a sequential framewok.

[4] The fixed effect has a different impact on birth spacing and mortality.

[5] Wolpin (1984) develops a finite-horizon dynamic stochastic model of discrete choice with respect to life-cycle fertility in a world where infant survival is uncertain and offers results for the number, timing and spacing of children for exogenous child mortality. We however choose to focus on Rosenzweig beucase it examines the effects of spacing on child survival.

2.1. Health Production Function

Rosenzweig (1986) applied the Beckerian framework to a three-period model to determine how birth spacing may affect birth outcomes. A key feature of this model is the health production function, which plays the same role as the child quality production function in the Becker model.

Assume that the quality H_{ij} (e.g., health) of a child *i* born to family *j* depends on its birth order *i*, age of its parents when born S_i, intervals between its birth and both prior and subsequent births and child specific resources Z_{ij}. Thus for a child of order *i* in family *j* (among n children) who is neither the first nor the last child (in linear form) we have:

$$H_{ij} = \gamma_a S_{ij} + \gamma_p (S_{ij} - S_{i-1j}) + \gamma_n (S_{i+1j} - S_{ij}) + \sum_{k=2}^{n-i} \gamma_{sk} (S_{i+kj} - S_{i+(k-1)j})$$

$$+ \gamma_h^i Z_{ij} + \delta_j + \varepsilon_{ij} \qquad 1.$$

where γ_a parental age effect, γ_p prior spacing effect, γ_n is the posterior spacing with immediately subsequent sibling while γ_{sk} is the posterior spacing with respect to all other subsequent siblings[6] and δ_j is a family quality component common to all members of family *j* and ε_{ij} is a child specific random component. For the first child (i=1), $\gamma_p = 0$ while for the last child ($i=n$) γ_{sk}=0.

Equation (1), describing the health production technology, displays jointness of spacing decisions – an increase in the parental age of birth of child i affects its older sibling's post-birth interval and its younger sibling's prior birth interval. This interdependence implies that, even in the absence of differences in child specific endowments (ε_{ij}=0), it is unlikely that parents could simultaneously equalize child specific resources Z_{ij} across children and equalise quality outcomes because of the sequential nature of childbearing. Thus child specific investments in health will be correlated with birth order and spacing as well as children's endowments.

In the context of competition among siblings for limited resources, we argue that prior and posterior birth spacing reflect the age difference between consecutive siblings and capture an important aspect of the intensity of competition among siblings that is little discussed in the literature.[7] Multiple birth (twins/triplets) too will naturally impose a strain

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[6] Note that in our empirical estimation, posterior spacing with respect to other subsequent siblings is never significant and that is why the rest of our analysis focuses on γ_n for the immediately next sibling.

[7] This works in conjunction with other possible factors, for example, cultural preference for sons in certain societies or the biological factors (e.g., maternal depletion due to shorter birth interval). See further discussion later in the section.

on parental resources and may thus increase the competition between the current siblings as well as that between current and existing siblings.[8]

The effects of siblings on child mortality would be further complicated if parents are not only resource constrained, but are also characterised by preferences for sons either because of the higher expected earnings of boys (Rosenzweig and Schultz, 1982) or, even, prejudice. We estimate the model separately for male and female children to allow for possibility that the treatment of a child may depend on its gender. Some researchers (such as Garg and Morduch, 1998 and Butcher and Case (1994)) have used, respectively, the number of female children and 'any daughters' to test for favourable treatment of males. We allow for this possibility within our gender specific estimating equations by including the gender of the first born as an instrument for number of girls.[9]

2.2. Data

India is an interesting case to consider in the present context. Child mortality rates for girls are among the highest in the world.[10] There is also an interesting regional variation within the country. 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.[11] Our sample is drawn from the eastern Indian state of West Bengal. In the post-independence period, West Bengal started its economic development in a relatively good position among the Indian states as reflected in its high rate of urbanisation, 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)[12] in rural West Bengal has declined between 1981 and 1990, the state's own rate of decline in

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[8] Problem of endogeneity is common in the demographic literature. Twins have often been used as an identifying strategy in the demographic literature. For example, see use of twins on first birth as in Rosenzweig and Wolpin (1984, 2000) or twins in later births as in Black et al. (2004). Here we focus on the effect that arrival of twins may have on child mortality.

[9] The number of sisters (or brothers) depends on the choice of family size and is therefore endogenous. Although the gender of a particular child is random, the probability of having a sister increases with the number of siblings. However, the gender of the first child cannot be correlated with the gender and other aspects of the second child although it is correlated with number of children of a particular gender and can therefore be used as an instrument.

[10] 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.

[11] Though the female mortality rates are generally lower in the Western countries.

[12] Number of infants dying before reaching one year of age, expressed per 1000 live births in a year.

the 80s 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). Table 1 compares West Bengal's demographic performance with important Indian states in 1991.

We use the National Family Health Survey (NFHS) 1992-93[13] household-level data from rural and urban West Bengal. This allows us to construct a complete birth history for each woman aged 13-49 years. Given that in our sample the death rate tails off from age five onwards, age is right censored at 60 months. There are 12,902 children in our sample of whom 51% are male. 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).

A preliminary analysis of the data (shown in Table 2) suggests that the mortality rate for children during their first 5 years is about 13% across the whole sample. It rises slightly when there is more than one child and birth spacing is less than 12 months but more than doubles when we consider non-first born children with birth intervals of a year or less. The mortality rates are even higher when the child is one of the twins or if the first child is a female. Gender differences are also observed in these estimates, though the extent is rather limited except when the first born is female. If the first born is a female and the birth spacing is a year or under, then subsequent females are over 30% more likely to die in the first 5 years. It however follows that mortality rates decline if the spacing between consecutive births is between 24-60 months though beyond 84 months the mortality rate may go up somewhat for the non-firstborn children in our sample.

2.3. Methodology

The unit of observation is a woman together with the birth history of all her children. The primary hypothesis is that child mortality depends on both prior and posterior birth spacing. Since we do not observe prior birth spacing for first born children and posterior

[13] The second NFHS undertaken in 1998-99 was designed to strengthen the database further and facilitate implementation and monitoring of population and health programmes in the country. Though some additional information (e.g., height and weight of all eligible women, blood test for women and children) were collected, the information that we use remained very similar. Our preliminary analysis also yielded similar results as reported here.

birth spacing for youngest children, we concentrate on middle order children.[14] We first model child mortality as a probit equation showing the probability of a child dying in the first 5 years of life. In an alternative specification, we also estimate a mortality hazard equation (see Appendix).

The household chooses the number and age composition (reflecting birth interval) of its children to maximise the present value of income produced by all family members. This income stream depends on the survival prospects of the children. The optimal values of different child variables, such as the number of children and birth spacing, will therefore depend in part on the values of the error term in the mortality equation. If this error term incorporates factors that are constant over children for the family but unobserved in the data, the values of any birth spacing variables in the mortality equation may be correlated with the error. We have attempted to resolve the resulting problems elsewhere using instrumental variables[15] but the use of weakly correlated instruments may actually exacerbate the problem. Here, we model the source of the endogeneity as a fixed effect reflecting unobserved family-specific heterogeneity that affects both mortality and birth spacing. We then introduce hazard equations to explain birth spacing. Assuming that the fixed effects in the birth spacing and mortality equations are correlated, we estimate the mortality equation purged of the correlation between its error and the birth spacing variables.[16]

The model is identified by its recursive structure and the covariance restrictions imposed by the inclusion of a fixed effect in each equation. This issue is discussed in Chamberlain and Griliches (1975).[17] The non-linear form of the model also guarantees identification.

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[14] We have also tried to include all children in our estimation. In this case, prior spacing for oldest child was estimated by the time between mother's age at marriage and birth of the first child while posterior spacing for the youngest child was the time elapsed between the birth of the child and the time of the survey (for non-sterilised couple) or the time the couple was sterilised. However the log-likelihood function would not converge probably because of the poor quality of the available information (age at marriage, number of marriages or time of sterilisation), which in turn resulted in rather sporadic distribution of prior/posterior spacing of the oldest/youngest children in our sample. Note that the estimation of prior and posterior spacing hazard equations indirectly takes account of first born and youngest children as censoring variables.

[15] Makepeace and Pal (2001).

[16] In practice, we estimate the model jointly by maximum likelihood. An analogue to this procedure is the treatment model using Heckman-type selection adjustments to correct the error for omitted variable bias. To pursue this analogy, the mortality equation models the outcome of the treatments (birth spacing) and the birth spacing equations the selection into the treatment.

[17] A short note on this is available from the authors on request.

Child mortality equation

The mortality equation shows the probability that a child dies within 5 years of birth. The propensity to die for the *i*-th child born to *j*-th mother is given by:

$$D_{ij}^* = \beta_1 Z_{Mj} + \beta_2 X_{Mij} + \beta_3 PREV_{ij} + \beta_4 NEXT_{ij} + \delta_j + u_{Mij} \qquad 2.$$

The child dies if $D_{ij}^*>0$ and death is recorded by the dummy variable, D_j, that takes the value 1 if the child has died. *PREV* and *NEXT* are the prior and posterior 'birth spacing' variables showing the lengths of time between the birth of the current child and the births, respectively, of the previous child and the next child. X_M and Z_M are respectively vectors of exogenous. child-specific and household-specific covariates. We adopt a probit specification that enables family-specific differences, δ_j, to be modelled as random effects. u_{Mij} is a random error independently and identically distributed with zero mean and unit variance.[18]

PREV and *NEXT* reflect the potential effect that sibling competition for limited parental resources has on child health outcomes since rivalry may decline as the age gap increases. Thus parents can devote more time and efforts to bringing up a child if either prior or posterior birth spacing is longer, especially since this will also involve less maternal depletion.

In general, the probability that the *i*-th child dies will depend on a vector of other characteristics. Among the individual child specific characteristics, we include if the current child is a twin. Multiple birth (twins/triplets) may impose a strain on parental resources and may thus increase the competition among the current siblings as well as that between current and existing siblings.

If parents are characterised by son preferences, the gender of the current child could be important determinant of child mortality. In order to reduce the effects of pro-male bias in the pooled sample, we separately estimate the mortality functions for boys and girls in our sample; this allows us to examine how the same set of individual/household characteristics may affect survival of male and female children differently.

All the remaining covariates are household-specific. Preferences for sons in the Indian society are found to be important in birth spacing and therefore in child survival. As explained earlier, we examine whether gender of the first child (Firstfem) has a direct/indirect (via its effect on spacing) impact on mortality. In addition, we include a number of variables reflecting various aspects of health. The dummy for 'whether the first

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[18] In an alternative specification, we also estimate a mortality hazard equation to compare with the mortality probit estimates. These hazard estimates are shown in Appendix Table A1.

child died' may take account of 'death clustering' such that families experiencing child death may have shorter birth intervals (Dasgupta, 1997) and higher mortality rates. We have also included a variable to indicate if the current child is one of twins.[19] The latter can be treated as another health variable since it will be associated with factors such as low birth weight (although competition for resources will also play a role).

The provision of public services like safe drinking water, sanitation or use of other health inputs (like immunisation) will also affect child health.[20] We did try to include both access to modern toilet and safe drinking water in the mortality equation, but none of them turn out to be significant. As an alternative, we tried including a binary variable for rural residential location because provision of public services tend to be worse in rural areas of the state. This rural dummy would account for the effects of inter-regional (rural/urban) variation in public services on child mortality within the state. Note however that residential location (rural/urban) is the location at the time of the survey and may not reflect location during first five years of a child's birth especially if the family moves over time.

Our model emphasised the role of family resources. Since literate mothers tend to be more educated[21] and from higher income families, we use mother's literacy as a proxy for income and wealth. Since NFHS data do not provide any information on household income or expenditure, we also include some key household assets variables, namely, ownership of land (Agland) and brick-build houses (Pucca), to control for variation in wealth effects. Religion may also be considered to be an important determinant of socio-cultural practices, e.g., defining pro-male bias, which in turn could affect parental allocation for investment in children. To this end, we include a dummy for Muslim children.[22]

In an alternative specification, a proportional hazard model of mortality is estimated (see Appendix). In addition to the above mentioned variables, in this case we include three

[19] 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.

[20] We however cannot analyse the effects of specific health inputs (e.g., prenatal care, hospital delivery or child vaccination) on child mortality (e.g., Maitra 2004) since these information were only collected for children born in the last 3 years (this holds for both rounds of NFHS).

[21] 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. Hence, we could not include comparable characteristics of the father as we did for the mother.

[22] Most households in our sample are Hindus or Muslims. There are only a minority of households belonging to other religions. We tried various specifications, but only Muslim dummy was turned out to be robust and consistently significant.

more variables to capture the baseline hazard. In particular, we define two nodes, namely, 3 and 6 months and using these two nodes we create three additional variables: if the child dies between 0-3 months, 3-6 months and above 6 months. Each new variable represents the original spacing variable on a specific segment of its range so that the estimated effect of the splines is no longer linear, but piece-wise linear. These spline coefficients may directly be interpreted as slope coefficients (Panis, 1994).

Birth spacing equations

Posterior spacing: The log hazard rate of spacing from the time of birth of child i (born to jth woman) till the arrival of the next sibling (*NEXT*) is a function of calendar time ($T(t)$) and household (Z_{2j}) and individual child-specific (X_{2jj}) characteristics and a family-specific[23] heterogeneity component ε_j common to all children in j-th family. It is:

$$\text{Ln } h_{ij}^{NEXT}(t, \varepsilon_j) = \alpha_0 + \alpha_1 Z_{Nj}(t) + \alpha_2 X_{Nij}(t) + \alpha_3 T(t) + \varepsilon_j + u_{Nij} \qquad \textbf{3.}$$

This model is proportional in the sense that the hazard is characterised in terms of a baseline hazard that captures duration dependency and proportional shifts of the baseline hazard..

Prior birth spacing: In a similar fashion, time since the birth of the previous sibling (*PREV*) is specified as follows :

$$\text{Ln } h_{ij}^{PREV(}t, \eta_j) = \gamma_0 + \gamma_1 Z_{Pj}(t) + \gamma_2 X_{Pij}(t) + \gamma_3 T(t) + \eta_j + u_{Pij} \qquad \textbf{4.}$$

The baseline hazard for each birth spacing equation is defined as piecewise linear splines which depends on two nodes. We have defined two nodes as 12 and 24 months as we find that the mortality risks are higher within the first two years of a child's life. Using these two nodes, we create three variables, namely, if spacing is 12 or less months, greater than 12 months but less than or equal to 24 months, and greater than 24 months.

Each birth spacing hazard equation depends on both individual and parental/ household characteristics, with some identifying variables between the two.[24] Among the variables present in both the spacing equations, we include mother's age at first birth[25] and

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23 The observations are grouped by mother so the factor is strictly speaking mother-specific. However, family break-ups are extremely rare so we interpret this more broadly as a family-specific effect.

24 See Appendix for the definition of these variables.

25 Without much loss of generality we could treat mother's age at first birth for the rural sample as an exogenous variable (evidence suggests that use of contraception is almost non-existent before the birth of the first child), though we could not ignore the element of simultaneity while including mother's age at each birth of the child.

mother's literacy.[26] Mother's age at first birth is a good measure of fecundity while mother's literacy is widely found to reduce fertility. We also include a binary variable indicating delivery problems in previous births, if any. This is likely to affect both spacing equations. Variation in household wealth is controlled by including ownership of land (Agland) and whether the household lives in a brick house (Pucca).

The choice and use of current contraceptives are important determinants of birth spacing in many cases, though, they are chosen by the couple in question and therefore, could not be treated as exogenous. Hence we use proxies that can reflect use of contraception in our sample. We use a binary variable indicating whether the couple in question belongs to a Muslim family. There is evidence that contraception use is rather limited among the Muslim couples in our sample, thus the binary variable Muslim could capture the couple's attitude towards modern contraception.[27] We use two other binary variables, namely, ownership of radio (radio) and television (tele), to indicate couple's awareness towards contraception through media advertisements. As with the mortality equation, we include the characteristics of the children already born, for example, whether the first child is a female and if the first child is dead.

Identification

Identification is achieved by the recursive structure of the model and the implied covariance restrictions implied by the correlated fixed effects. In a sequential framework, prior spacing is important for the posterior spacing decision and mortality of the context child and not vice versa. This is because timing of these two spacing decisions are separated by the birth of the context child. Once the child is born, parents can only move forward to plan for posterior spacing, taking account of prior spacing experience (and cannot go backward in time to revise the prior spacing decision already taken before the birth of the context child).

In addition, there are some differences in the lists of other regressors between the equations. Whether the current child is one of the twins (Twin) is important for posterior spacing decision, but not for the prior spacing of the context child (because parents cannot

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[26] This is the first principal component of all different assets variables the household may own.

[27] Analysis of NFHS 92-93 data (see Table 1A) suggests that compared to Hindus, a significantly larger proportion of Muslim couples use no contraception. This difference is mostly accounted for by the difference in the proportion of sterilised Muslim couples; in particular, compared to Muslim couples, more than double of the Hindu couples are sterilised. In contrast, use of modern (use of pills, IUD/copper, injections and condoms)or traditional (abstinence/withdrawal) non-terminal methods of contraception is rather comparable among various religious groups in our sample..

know before conception whether the child is a twin). Hence, Twin is included only in the posterior spacing equation. Similarly there are also some variables identifying the mortality equation. Ownership of radio and television is included only in the spacing equations as indicators of parental awareness of modern contraception. While some may argue that these are also household assets and could go in all three equations. But the fact remains that ownership of brick-build house and land would proxy for the ownership of any other valuable assets in the model. Secondly, the binary variable Prevprob indicating delivery problems at previous births is included in the spacing equations though not in the mortality equation because we believe that this variable is more likely to affect child health indirectly through spacing. If, however, there is a genetic problem attached to delivery problem of a particular couple and if it repeats itself for each birth, this would enter the mortality equation through the unobserved mother-specific heterogeneity factor that we explain below.

Parent-specific unobserved heterogeneity

Since decisions on both birth interval (prior and/or posterior) and investment for child survival are made by the same woman (or couple), the residuals are likely to be correlated across decisions. We therefore have two components in each residual: a mother/family specific (η, ε, δ) component and a child specific (u_N, u_P, u_M) component respectively in the posterior and prior birth spacing and the mortality equations. The family-specific components are constant across all births of a given mother. Each is assumed to be distributed normally with zero means and variances σ^2_η , σ^2_ε and σ^2_δ respectively. The child-specific components are normally, independently and identically distributed with unit variance, and independent of the family specific components. The correlation coefficients between the aggregate errors in the different equations are shown by ρ_{KL} where $K,L=N,P,M$ respectively for posterior, prior spacing and mortality equations.

Joint Estimation

Joint estimation of the spacing hazard and the child mortality probit equations is based on maximization of the joint marginal likelihood function obtained by integrating the product of conditional likelihood functions over the range of unobservables, weighted by the joint density function of unobservables.[28] The conditional likelihoods are the probabilities of observed outcomes (the birth spacing hazard and child mortality probit equation for each child in the sample), conditional on the vector of unobserved heterogeneity components

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28 The estimation is based on the technique followed by Panis and Lillard (1994, 1995).

($\eta, \varepsilon, \delta$). Thus when prior (PREV) and posterior (NEXT) spacing are treated as endogenous in the child mortality probit regression, the joint marginal likelihood function is written as:

$$\int_{\varepsilon_i}\int_{\eta_i}\int_{\delta_i}\left[\prod L^n(\varepsilon_i)\prod L^P(\eta_i)\prod L^S(\delta_i)\right]f(\varepsilon_i,\eta_i,\delta_i)\,d\varepsilon\,d\eta\,\,d\delta) \qquad 5.$$

where $f(\varepsilon, \eta, \delta)$ is the joint distribution of the unobserved heterogeneity components in the three equations and is a three dimensional normal distribution characterised as follows:

$$\begin{pmatrix}\varepsilon_i\\ \eta_i\\ \delta_i\end{pmatrix} \sim N\left(\begin{pmatrix}0\\0\\0\end{pmatrix},\begin{pmatrix}\sigma_\varepsilon^2 & & \\ \rho_{\varepsilon\eta}\sigma_\varepsilon\sigma_\eta & \sigma_\eta^2 & \\ \rho_{\varepsilon\delta}\sigma_\varepsilon\sigma_\delta & \rho_{\eta\delta}\sigma_\eta\sigma_\delta & \sigma_\delta^2\end{pmatrix}\right) \qquad 6.$$

The model is estimated using Full Information Maximum Likelihood (FIML).

3. RESULTS

Tables 4A, 4B and 4C report separate estimates by gender of the child mortality (probit) equation and the spacing (hazard) equations. For completeness, we include 'base-line' estimates that ignore the family specific effect ('the no-heterogeneity' results). For each gender, we also present estimates assuming alternatively that the family-specific effects are uncorrelated and correlated. The same factors are significant in both uncorrelated and correlated estimates so the results are not qualitatively sensitive to this assumption (though these results differ somewhat between male and female children). The magnitudes of most of the estimates are approximately the same in each set of results while the value of the log likelihood function is higher for the correlated estimates.

The cross-correlations between the errors in the hazards and the mortality equation in Table 4A are highly significant for males and significant for females so we concentrate on these correlated results. Later, we shall demonstrate that the uncorrelated estimates can underestimate the probability of death. The negative values of the correlation coefficients suggest that unobserved factors that increase the instantaneous chance of either type of spacing (i.e. shorten either the time to the next birth or time since the last birth) simultaneously tend to lower the chance of a child dying. This is consistent with our basic hypothesis that the smaller the interval between births the lower (higher) the chances of survival (mortality).

Estimates of Child Mortality

Table 5B presents the estimates of the child mortality equation. The correlated estimates of mortality confirm our central hypotheses that an increase in the length of time either since the birth of the previous child (*PREV1*) or to the birth of the next child (*NEXT1*) lowers the chance of the child dying in the first 5 years of life. Similar results are obtained from the alternative mortality hazard specification, which are summarised in Appendix Table A1. Secondly, being one of the twins increases the risks of mortality for both male and female children in our sample, again emphasizing the aspect competition for limited resources both inside and outside the mother's womb. Death of the first child too increases the mortality risks of subsequent female children, perhaps suggesting some pro-male bias in response to this kind of tragedy. Pro-male bias is evident in other respects as well. For example, boys (and not girls) from families where the first child is a female, are more likely to survive. Thus sibling age and gender composition plays a central role in explaining child mortality. Role of mother's education is confirmed since both male and female children of literate mothers are significantly less likely to die. Among the assets variable, boys living in brick houses are more likely to survive though none of the asset variables are significant for the girls. Religion may also be important as Muslim boys are significantly less likely to die, again suggesting some pro-male bias prevalent in the Muslim community.

After controlling for all other factors, living in a rural region has however no significant effect on child mortality. Even if individuals living in rural areas are less well-off and have poorer access to public health facilities, they do not fare worse than those in towns and cities. The latter may be facilitated by the availability of certain other types of public goods like cleaner air so that the net effect of living in a rural region is not necessarily negative. Or it may simply reflect the fact the rural location at the time of the survey cannot simply capture the effect of location at the time of the birth of the child.

Estimates of Birth Spacing

The 'baseline' hazard of having a subsequent sibling is greatest in the first 12 months. It then declines gradually from 12-24 months and then after 24 months (note that the coefficients of DURSP1, DURSP2 and DURSP3 gradually decline). Among various socio-economic variables, some tend to affect prior and posterior spacing differently. For example, posterior hazard is lower though prior hazard is higher for boys born to literate mothers. Similar asymmetric wealth effect is noted among households living in a brick house: the posterior hazard is lower and prior hazard higher for more wealthy households living in a brick house, though these effects are significant for male children only. Mother's age at first birth is however important for both boys and girls. Boys and girls born to older mothers tend to have lower posterior hazard. Previous delivery problems of the mother however significantly lower the posterior hazard of female children, but is not important for boys. As with mortality, household religion tends to be more important for girls: Muslim girls are more likely to face a higher prior hazard than other girls. Regional location (e.g., rural) however remains insignificant in the spacing equation.

Among the sibling composition variables, the hazard of having a subsequent sibling is higher if the first child is a female and the effect is significant only for the male children. Death of the first child however significantly shortens the prior spacing while longer prior spacing significantly lower posterior hazard for both male and female children.

3.3. Inferences

Thus these correlated estimates of birth spacing and child survival generally lend support to the central hypothesis of sibling rivalry in that shorter birth interval (prior and posterior) and twin births significantly enhance mortality risks among 0-5 year old male and female children. In general the parameter estimates from uncorrelated[29] and correlated models indicate similar pattern of results though uncorrelated estimates are likely to suffer from endogeneity bias. In order to understand the extent of the bias in the uncorrelated estimates, we finally compare the predicted probability of mortality for the middle order children, as summarised in Table 5. These predicted probability estimates not only suggest a significant higher mortality risks if consecutive children are born within 12 months and if the current child is one of the twins, but also that the uncorrelated estimates tend to underestimate[30] the mortality risks in our sample. It also highlights the asymmetry between prior and posterior spacing between male and female children in our sample. In particular, if the context child is a male, shorter prior/posterior spacing does not make much difference in the mortality risks (the risk is only slightly higher if the prior spacing is shorter). If however the context child is a female, mortality risks are substantially higher if the posterior spacing is less than a year than if the prior spacing is less than a year. These estimates further

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[29] where birth intervals are treated as pure exogenous variables in the mortality equation.

[30] This is the net effect of allowing for unobserved heterogeneity with non-zero correlation. Note that coefficients of variances ($\sigma_{\eta}, \sigma_{\varepsilon}, \sigma_{\delta}$) are positive while two of the three correlation coefficients ($\rho\,(\eta\delta)$, $\rho\,(\varepsilon\delta)$) are negative.

σ_{δ}

$\rho\,(\eta\varepsilon)$

$\rho\,(\eta\delta)$

$\rho\,(\varepsilon\delta)$

substantiate the role of siblings on child mortality in resource constrained households with pro-male bias.

4. CONCLUDING COMMENTS

This paper examines the role of siblings on child survival in India and argues that competition among siblings for limited resources plays a significant role in child survival. Within a sequential framework, this means that, even if child-specific unobserved endowments are identical, an increase in parental age at birth of the child affects its older sibling's posterior spacing and younger sibling's prior spacing. This interdependence means that it is unlikely for parents to simultaneously equalise child specific resources across siblings and thus the quality outcomes. Parental allocation of resources is further complicated if parents are characterised by pro-male bias. Thus in addition to the competition between twins, age (measured by prior and posterior spacing) and gender composition (measured by gender of the current as well as first child) become important for child survival.

The empirical analysis based on the recent NFHS data from West Bengal employs a likelihood estimation technique to determine birth spacing hazard and mortality probit equations as correlated processes, allowing for mother/parents specific unobserved heterogeneity among male and female children. The explanatory variables are chosen to reduce the endogeneity bias as far as practicable and include among various individual and household specific characteristics, prior and posterior birth spacing, if the child is one of the twin and also gender of the first child respectively as measures of age and gender composition of siblings. These devices allow us to re-estimate the effects of sibling composition, corrected for the possible endogeneity bias that has not previously been attempted in this literature.

Given the values of other variables, we interpret our results as showing that competition for limited resources is an important part of any explanation of child mortality in West Bengal. Direct sibling rivalry is captured by the prior and posterior birth spacing. As the birth spacing increases, the chances of survival improve for the context child perhaps because parents are able to devote more time and effort to bringing that child through his or her critical early years. Twin birth too significantly enhances the mortality risks of both male and female children while risk of having a subsequent sibling is higher for boys if the first child is a female.

It has widely been documented that mother's literacy and household assets may affect birth interval. A particular advantage of our modelling strategy is that these sequential estimates allow us to establish how these factors may affect prior and posterior spacing differently. We show that mother's literacy and household assets may lower the hazard of subsequent birth, but may still increase the hazard of prior birth. The essential implication is that these variables may be more effective to reduce fertility once a target family size is achieved.

Predicted probability estimates substantiate the bias generated by ignoring the possible correlation between the two decisions. These estimates also clarify the role of siblings on risks of child mortality. In general there is evidence that prior and posterior birth intervals have different effects on young boys and girls, which in turn reflect the nature of decisions made by resource constrained parents characterised by promale bias in the Indian state.

There is thus a significant potential for reducing child mortality even in a state like West Bengal (with moderate level of female literacy among the Indian states) and this could be achieved by encouraging use of modern non-terminal methods of contraception for spacing birth. The potential effects of reducing child mortality by spacing child birth could be far more in some other Indian states with lower levels of female literacy.

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Table 1A. Comparison of West Bengal with important Indian states

States	Popln. (in mn) 1991	Female literacy Age 7+ 1991	Female labour participn 1991	Total fertility rate	Infant MortalityRate 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.nic.in/mohfw/popindi.html

Table 1B. Current contraception use among various religious groups

	% of the particular religious group		
Current method of contraception	**Hindu**	**Muslim**	**Other**
None	**44**	**60**	**47**
Traditional non-terminal (abstinence/withdrawal)	**18**	**19**	**21**
Modern non-terminal (pills, condoms, etc)	**6**	**6**	**2**
Sterilisation (male & female)	**32**	**15**	**30**

Table 2. Effects of sibling composition on child mortality

(Percentages in categories)

Birth interval months	Birth order	Gender of first born	First born died	Male	Female	All
All children	All	All	No	13.3	13.2	13.2
12 or less	All	All	No	16.4	14.5	15.5
12 or less	Not first	All	No	28.8	29	29
12 or less	Not first	Female	No	27.3	35.7	31.3
12 or less	Not first	All	Yes	34.3	34.6	34.5
>12 & <=24	All	All	No	12.6	13.6	13.1
>12 & <=24	Not first	All	No	18.8	20.3	19.5
>12 & <=24	Not first	Female	No	16.1	21.8	19.0
>12 & <=24	Not first	All	Yes	21.8	23.4	22.6
>24& <=60	All	All	No	8.0	7.0	7.5
>24& <=60	Not first	All	No	12.4	10.9	11.7
>24& <=60	Not first	Female	No	12.5	10.0	11.2
>24& <=60	Not first	All	Yes	15.9	16.2	16.1

Table 3: Sample characteristics - means and standard deviations

name		Female Mean	Female Std Dev		Male Mean	Male Std Dev
Delivery problem in previous births (PREVPROB)	3067	0.06195	0.241104	3044	0.050591	0.219198
Mother's age at birth (AGEMUM1)	3067	17.14803	2.630718	3044	17.06439	2.671485
Mother is literate (LITMUM)	3067	0.297033	0.457026	3044	0.290079	0.453873
Child is one of the twins/triplets (TWIN)	3067	0.017933	0.132729	3044	0.020368	0.141279
First child is a female (FSTFEM)	3067	0.515814	0.499831	3044	0.494087	0.500047
First child is dead (FSTDIE)	3067	0.257255	0.437192	3044	0.268068	0.443026
Time since the birth of the previous child (PREV1)	3067	30.59374	14.84545	3044	30.5138	15.30044
Time to the birth of the next child (NEXT1)						
Ownership of land (AGLAND)	3067	0.546136	0.497948	3044	0.542707	0.498255
Ownership of brick house (PUCCA)	3067	0.108901	0.311566	3044	0.112681	0.316254
Ownership of radio (RADIO)	3067	0.361265	0.480446	3044	0.361367	0.480475
Ownership of television (TELE)	3067	0.084773	0.27859	3044	0.086071	0.280515
Muslim household (MUSLIM)	3067	0.34431	0.47522	3044	0.367608	0.482233
Lives in rural region (RURAL)	3067	0.846104	0.360908	3044	0.832457	0.373521

Table 4A. Structure of unobserved heterogeneity

	Uncorrelated		Correlated	
	Male	Female	Male	Female
σ_η	0.2615 ***	0.2591 ***	0.5330 ***	0.4231 ***
	0.0657	0.0583	0.05	0.0457
σ_ε	0.6242 ***	0.5569 ***	0.5944 ***	0.4661 ***
	0.0382	0.0393	0.0424	0.0409
σ_δ	0.3447 ***	0.4248 ***	0.3390 ***	0.3713 ***
	0.0961	0.0706	0.1055	0.0801
ρ ($\eta\varepsilon$)			0.8953 ***	0.8941 ***
			0.0633	0.1331
ρ ($\eta\delta$)			-0.2221 **	-0.2129 *
			0.0822	0.1278
ρ ($\varepsilon\delta$)			-0.2246 ***	-0.3812 *
			0.0508	0.2216

Table 4B. Uncorrelated and Correlated Estimates of Mortality

	Uncorrelated Survival estimates				Correlated estimates	
	Male		**Female**		**Male**	**Female**
	No het	Het	No het	Het	With het	With het
Intercept	-0.0291	-0.0279	-0.2734 **	-0.3064 **	0.02	-0.5666 ***
	0.1143	0.1269	0.1106	0.1306	0.1911	0.1658
LITMUM	-0.1685 **	-0.1773 **	-0.2283 ***	-0.2544 ***	-0.1674 **	-0.2510 ***
	0.067	0.0754	0.0634	0.0766	0.0762	0.0754
TWIN	1.2387 ***	1.2886 ***	1.1822 ***	1.2583 ***	1.2551 ***	1.3386 ***
	0.1702	0.1901	0.154	0.1801	0.1969	0.1822
FSTFEM	-0.0940 *	-0.101*	-0.033	-0.0408	-0.1013*	-0.0525
	0.0549	0.0608	0.0539	0.0653	0.054	0.0642
FSTDIE	0.0853	0.0961	0.1446 **	0.1608 **	0.1066	0.1714 **
	0.0593	0.0683	0.0604	0.0741	0.0695	0.0728
PREV1	-0.0182 ***	-0.0193 ***	-0.0155 ***	-0.0165 ***	-0.0195 ***	-0.0105 ***
	0.0021	0.0023	0.0021	0.0023	0.0034	0.0032
NEXT1	-0.0113 ***	-0.0118 ***	-0.0102 ***	-0.0110 ***	-0.0127 ***	-0.0077 ***
	0.0017	0.0018	0.0014	0.0015	0.0025	0.0022
AGLAND	-0.0281	-0.0277	-0.0044	0.0019	-0.0288	-0.0013
	0.0576	0.066	0.0572	0.0695	0.0668	0.0678
PUCCA	-0.1735 *	-0.1846*	-0.109	-0.105	-0.1954 *	-0.1445
	0.1019	0.1033	0.1025	0.1213	0.1164	0.1185
MUSLIM	-0.2026 ***	-0.2212 ***	-0.0285	-0.036	-0.2082 ***	-0.0099
	0.057	0.0659	0.0571	0.0696	0.0676	0.0684
RURAL	0.0532	0.0395	0.0812	0.085	0.038	0.0717
	0.0843	0.0941	0.0819	0.0984	0.095	0.0966

Note: Standard errors are shown below. Levels of significance: *- 10%; **- 5%; ***- 1%

Table 4C. Correlated and uncorrelated estimates of prior and posterior spacing, Male

	Uncorrelated estimates				Correlated estimates	
	Spacing Posterior (*NEXT*)		Spacing Prior (*PREV*)		Posterior (*NEXT*)	Prior (*PREV*)
	No het	With het	No het	With het	With het	With het
0-12 months	0.6815 ***	0.6720 ***	0.6794 ***	0.6924 ***	0.6817 ***	0.6701 ***
	0.0955	0.0956	0.0854	0.086	0.1069	0.0936
12-24 months	0.1494 ***	0.1535 ***	0.1263 ***	0.1480 ***	0.1587 ***	0.1445 ***
	0.0063	0.0066	0.0057	0.0061	0.0076	0.0068
> 24 months	-0.0012	0.0022	-0.0033 **	0.0137 ***	0.0070 ***	0.0114 ***
	0.0011	0.0019	0.0014	0.0018	0.0018	0.0021
Intercept	-12.1042 ***	-12.0419 ***	-12.3027 ***	-12.7032 ***	-12.6949 ***	-12.4216 ***
	1.1193	1.1212	1.0009	1.0155	1.2682	1.1066
PREVPROB	-0.1355	-0.1443	0.1	0.1254	-0.1021	0.0748
	0.0875	0.0956	0.0859	0.1363	0.1219	0.1275
AGEMUM1	-0.0139 **	-0.0148 **	-0.003	0.0003	-0.0153 *	0.0006
	0.0068	0.0075	0.0057	0.0089	0.0089	0.0087
LITMUM	-0.0877 **	-0.0931 *	0.0694 *	0.0936	-0.083*	0.0683*
	0.0445	0.0496	0.0409	0.0644	0.0417	0.04039
TWIN	0.0685	0.0804			0.0166	
	0.1507	0.1631			0.194	
FSTFEM	0.0731 **	0.0796 *	0.0384	0.042	0.0950 **	0.0309
	0.0359	0.0409	0.0337	0.0525	0.0484	0.0511
FSTDIE	0.0056	0.0093	0.1153 ***	0.1654 ***	0.0356	0.1535 ***
	0.0419	0.0476	0.0369	0.0581	0.0569	0.0567
PREV1	-0.0119 ***	-0.0116 ***			-0.0038 *	
	0.0014	0.0014			0.002	
AGLAND	0.0035	0.0043	0.0248	0.0291	0.0277	0.0464
	0.0381	0.0426	0.0354	0.0539	0.0514	0.0527
PUCCA	-0.1125 *	-0.1217 *	0.0966 *	0.1174	-0.1237*	0.1276*
	0.0629	0.0708	0.0585	0.0915	0.0699	0.0686

RADIO	-0.0067	-0.0059	-0.0236	-0.0294	0.0161	0.0083
	0.042	0.0465	0.0362	0.0573	0.0551	0.0552
TELE	-0.0542	-0.0429	0.0235	0.0057	-0.0315	-0.0404
	0.0693	0.0781	0.0734	0.1147	0.0979	0.112
MUSLIM	0.0446	0.0442	0.0478	0.0668	0.0534	0.0416
	0.0374	0.042	0.0336	0.0548	0.051	0.0534
RURAL	0.0349	0.0334	0.0221	-0.1122	-0.0244	-0.0934
	0.0532	0.0595	0.0492	0.0779	0.0722	0.0764
Ln-L	-9998.88	-9996.85	-9792.72	-9745.76	-20901.5	-20901.52

Table 4D. Correlated and uncorrelated estimates of prior and posterior spacing, Female

	Uncorrelated estimates				Correlated estimates	
	Spacing				Posterior	Prior
	Posterior (*NEXT*)		Prior (*PREV*)		(*NEXT*)	(*PREV*)
	No het	With het	No het	With het	With het	With het
0-12 months	0.7395 ***	0.7319 ***	0.8151 ***	0.8017 ***	0.7166 ***	0.8129 ***
	0.1071	0.1071	0.0861	0.0861	0.1237	0.1011
12-24 months	0.1360 ***	0.1400 ***	0.1232 ***	0.1424 ***	0.1471 ***	0.1438 ***
	0.0057	0.0059	0.0054	0.006	0.0072	0.0069
> 24 months	-0.0049 ***	-0.0014	-0.001	0.0129 ***	-0.0017	0.0106 ***
	0.0011	0.0017	0.0014	0.0017	0.0019	0.002
Intercept	-12.5171 ***	-12.4868 ***	-13.6931 ***	-13.7389 ***	-12.7705 ***	-13.7023 ***
	1.2603	1.2615	1.0091	1.0138	1.4678	1.1893
PREVPROB	-0.1379 *	-0.1448 *	0.0649	0.1112	-0.1457*	0.0462
	0.0708	0.0797	0.0681	0.1085	0.0886	0.0992
AGEMUM1	-0.0163 **	-0.0166 **	-0.0110 *	-0.0119	-0.0202 **	-0.0251 ***
	0.0069	0.0077	0.0059	0.0088	0.0092	0.0081
LITMUM	0.049	0.0513	0.0165	0.018	0.0991	0.0211
	0.0465	0.052	0.0402	0.0593	0.0609	0.0538
TWIN	0.0925	0.1365			0.1793	
	0.1087	0.1132			0.1255	
FSTFEM	-0.0055	-0.0069	0.0423	0.0455	-0.0138	0.0468
	0.0356	0.04	0.0333	0.0491	0.0474	0.0462
FSTDIE	-0.0517	-0.0611	0.0960 ***	0.1316 **	-0.0588	0.1005 *
	0.041	0.0462	0.037	0.0556	0.0555	0.0518
PREV1	-0.0091 ***	-0.0085 ***			-0.0059 ***	
	0.0013	0.0014			0.0019	

AGLAND	0.0072	0.0047	0.0086	0.0045	-0.0317	-0.0295
	-0.0383	-0.0432	-0.036	-0.0539	-0.0516	-0.0499
PUCCA	-0.0815	-0.0857	-0.0228	-0.0179	-0.0951	-0.0558
	-0.0692	-0.0756	-0.0583	-0.0847	-0.0894	-0.0778
RADIO	-0.064	-0.07	-0.0334	-0.0409	-0.0765	0.0062
	-0.0415	-0.0462	-0.0365	-0.0539	-0.0538	-0.0505
TELE	-0.0898	-0.0924	-0.0016	0.0056	-0.0746	0.0081
	-0.0816	-0.0896	-0.0776	-0.1094	-0.1035	-0.1011
MUSLIM	0.043	0.0537	0.0795 **	0.0847	0.0362	0.0828 *
	-0.0374	-0.0421	-0.0344	-0.0519	-0.0504	-0.0481
RURAL	-0.0606	-0.0654	-0.0812	-0.0828	-0.028	-0.0152
	-0.0551	-0.0609	-0.0506	-0.0734	-0.0721	-0.0695
ln-L	-10060.98	-10058.71	-9840.91	-9804.18	-21041.3	-21041.34

Table 5: Predicted probability of child mortality for middle-order children

	Uncorrelated	***Correlated***
All children with mean characteristics		
Male	0.127628	0.132122
Female	0.125070	0.134022
If prior birth interval <=12 months		
Male	0.22	0.25
Female	0.12	0.16
If posterior birth interval <=12 months		
Male	0.21	0.24
Female	0.67	0.69
If the child is a twin		
Male	0.59	0.69
Female	0.39	0.41

APPENDIX

Variable Definitions

The data are taken from the National Family Health Survey (NFHS) 1992-93 household 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
TWIN :	1 if the child is a twin or a triplet and 0 otherwise
FSTFEM :	1 if the first sibling in the family is a female and 0 otherwise
FSTDIE :	1 if the first sibling in the family died and 0 otherwise
PREV1 :	Length of time (in months) since the birth of the previous child
NEXT1 :	Length of time (in months) to the birth of the next child
PREVPROB :	1 if had some delivery problem in the previous birth
RADIO :	1 if the household owns a radio and 0 otherwise
TELE :	1 if the household owns a television and 0 otherwise
AGLAND	1 if owns land
PUCCA	1 if lives in a brick house
MUSLIM	1 if the family is Muslim and 0 otherwise
RURAL :	1 if the child lives in rural areas and 0 otherwise
MALE :	1 if the child is male and 0 otherwise

Table A1. Mortality hazard estimates

	Uncorrelated estimates				Correlated estimates	
	Male		Female		Male	Female
	No het	With Het	No het	With Het	With Het	With Het
DUR03	-1.3627 ***	-1.3459 ***	-1.4946 ***	-1.4641 ***	-1.3667 ***	-1.4872 ***
	-0.0835	-0.0852	-0.0953	-0.0951	-0.0868	-0.0989
DUR36	0.1194	0.1185	0.2980 ***	0.2932 ***	0.13	0.3201 ***
	-0.0811	-0.0811	-0.0887	-0.0881	-0.0827	-0.0925
DUR6+	-0.0404 ***	-0.0399 ***	-0.0355 ***	-0.0346 ***	-0.0393 ***	-0.0360 ***
	-0.0048	-0.0048	-0.0045	-0.0045	-0.0049	-0.0047
Intercept	-0.2558	-0.3720 *	-0.7796 ***	-0.9128 ***	-0.099	-0.8740 ***
	-0.1932	-0.2174	-0.1946	-0.2302	-0.2691	-0.3137
LITMUM	-0.3012 ***	-0.2961 **	-0.3906 ***	-0.4005 ***	-0.3311 ***	-0.3921 ***
	-0.1115	-0.1249	-0.1084	-0.1251	-0.1232	-0.1241
TWIN	1.8285 ***	2.0030 ***	1.6669 ***	1.8768 ***	1.9804 ***	1.8607 ***
	-0.1722	-0.2097	-0.1531	-0.2062	-0.2117	-0.2042
FSTFEM	-0.143	-0.1549	-0.068	-0.0652	-0.1753 *	-0.0502
	-0.0887	-0.1031	-0.0858	-0.1051	-0.1023	-0.1057
FSTDIE	0.1554 *	0.1793 *	0.2763 ***	0.2718 **	0.1605	0.2599 **
	-0.0925	-0.1086	-0.0968	-0.1194	-0.1082	-0.119
PREV1	-0.0332 ***	-0.0341 ***	-0.0267 ***	-0.0280 ***	-0.0383 ***	-0.0270 ***
	-0.0036	-0.0038	-0.0038	-0.004	-0.0053	-0.0054
NEXT1	-0.0227 ***	-0.0231 ***	-0.0207 ***	-0.0218 ***	-0.0259 ***	-0.0219 ***
	-0.0032	-0.0033	-0.0028	-0.0029	-0.0042	-0.0037
AGLAND	-0.0306	-0.0432	-0.0022	-0.0093	-0.0314	-0.0189
	-0.0916	-0.1076	-0.091	-0.1118	-0.1063	-0.1104
PUCCA	-0.2508	-0.2503	-0.1366	-0.1514	-0.2863	-0.1524
	-0.1617	-0.1817	-0.17	-0.1967	-0.1805	-0.1958
MUSLIM	-0.3757 ***	-0.3901 ***	-0.0986	-0.1019	-0.4006 ***	-0.0901
	-0.0922	-0.108	-0.0911	-0.1119	-0.1085	-0.1103
RURAL	0.0951	0.0786	0.1743	0.1401	0.0344	0.1231
	-0.1351	-0.1514	-0.1279	-0.1551	-0.1501	-0.1562
ln-L	-2355.14	-2350.86	-2494.52	-2486.32	-22028.9	-22229.9

Table A2. Mortality probit estimates with discrete spacing variables

	Uncorrelated estimates				Correlated estimates	
	Male		Female		Male	Female
Intercept	0.4579 **	0.4704 **	-0.0363	-0.0105	0.3891	-0.2873
	-0.1851	-0.1992	-0.1887	-0.2223	-0.2443	-0.2302
LITMUM	-0.1713 **	-0.1708 **	-0.2074 ***	-0.2419 ***	-0.1670 **	-0.2393 ***
	-0.0685	-0.0754	-0.0656	-0.0804	-0.0734	-0.0786
TWIN	1.2087 ***	1.2512 ***	1.1254 ***	1.2253 ***	1.2282 ***	1.3303 ***
	-0.1683	-0.1846	-0.1444	-0.1754	-0.1929	-0.177
FSTFEM	-0.0938 *	-0.0961	-0.0309	-0.0367	-0.0827	-0.0567
	-0.0557	-0.0622	-0.0538	-0.0671	-0.0614	-0.0659
FSTDIE	0.0899	0.0898	0.1497 **	0.1626 **	0.1297 *	0.1824 **
	-0.0603	-0.0676	-0.06	-0.0759	-0.0667	-0.0746
PREV2	-0.2978 **	-0.3034 **	-0.0452	-0.0522	-0.3158 **	0.0199
	-0.1319	-0.1386	-0.1444	-0.1636	-0.1409	-0.1599
PREV3	-0.6341 ***	-0.6595 ***	-0.3606 **	-0.4087 **	-0.5642 ***	-0.236
	-0.1274	-0.1343	-0.1448	-0.1634	-0.1455	-0.1646
PREV4	-1.0967 ***	-1.1962 ***	-0.5383 ***	-0.6208 ***	-0.8931 ***	-0.0862
	-0.2417	-0.2589	-0.2086	-0.2343	-0.2997	-0.2644
NEXT2	-0.7768 ***	-0.8012 ***	-0.6136 ***	-0.6998 ***	-0.8087 ***	-0.5451 ***
	-0.1404	-0.1484	-0.1214	-0.1353	-0.1531	-0.1392
NEXT3	-1.0101 ***	-1.0496 ***	-0.9495 ***	-1.0722 ***	-1.0530 ***	-0.8824 ***
	-0.1374	-0.1457	-0.1212	-0.1392	-0.1568	-0.1462
NEXT4	-1.1095 ***	-1.1476 ***	-1.0729 ***	-1.2099 ***	-1.1321 ***	-0.9157 ***
	-0.1835	-0.1946	-0.1677	-0.1863	-0.2208	-0.2106
AGLAND	-0.0329	-0.033	-0.007	0.003	-0.0468	-0.0059
	-0.0591	-0.0661	-0.0579	-0.0718	-0.065	-0.07
PUCCA	-0.1714 *	-0.1861 *	-0.1127	-0.0957	-0.1703	-0.1459
	-0.1032	-0.1127	-0.1043	-0.1259	-0.1127	-0.1225

MUSLIM	-0.1849 ***	-0.2029 ***	-0.055	-0.0717	-0.1532 **	-0.0516
	-0.0579	-0.0653	-0.0567	-0.0715	-0.0641	-0.0698
RURAL	0.0692	0.0611	0.1028	0.1122	0.0585	0.0842
	-0.0861	-0.0942	-0.0809	-0.1002	-0.0924	-0.0986
ln-L	-1236.31	-1234.46	-1252.36	-1244.42	-20929.25	21033.89