Understanding the Effects of Sibling Composition on Child Mortality: Evidence from India

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Abstract: This paper argues that spacing between consecutive births is an important aspect of competition among siblings for survival. Since parents simultaneously choose their desired values of birth spacing and the amount of time and other resources invested in children (which in turn affect child mortality), we use a maximum likelihood method to model birth spacing and child mortality as correlated processes while also allowing for family specific unobserved heterogeneity. Our estimates show that the chances of survival in the eastern Indian state of West Bengal increase with an increase in birth interval (prior and/or posterior) and decrease with the birth of a twin. Gender difference in child survival largely depends on religion and regional location, signifying the importance of socio-cultural practices in the state.

JEL Classification : D13, I12, O15

Key Words: Sibling competition, Birth spacing, Child mortality, Gender differences, Unobserved heterogeneity, Endogeneity bias.

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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 they need to allocate the 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 family structure (measured by birth intervals, presence of twins and gender of the first child) becomes crucial for determining child survival. Garg and Morduch (1998) quantified sibling competition 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. We thus argue that a better measure of sibling rivalry would be the age composition of siblings, which we measure by prior and posterior birth spacing within a sequential framework.

The Beckerian model (1991) explains the nature of parental investment in children 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), children with more sisters will be better off than children with more brothers. However, the Beckerian school of thought is essentially static in nature and therefore does not take account of the sequential nature of childbirth. An important exception in this respect is the study by Rosenzweig (1986) who studied the determination of optimum birth spacing in a sequential framework. Following Rosenzweig (1986), the present paper illustrates empirically the role of sibling age and sex composition on child survival.

Our analysis is based on complete birth history data obtained from the 1992-93 National Family Health Survey (NFHS) from West Bengal. We consider the birth history of women aged 13 to 49 years and find that the gender of the first child and the age composition of siblings (represented by birth spacing) are important indicators of sibling competition among children of different birth orders. Our results suggest that child mortality falls when the interval between births increases. Demonstrating this simple proposition raises a number of estimation problems that we attempt to address.

Firstly, parents simultaneously choose their desired values for birth spacing, the amount of time and other resources that are invested in children, which in turn affect child mortality. In this paper, we partly address the resulting estimation problems by modelling birth spacing and child mortality jointly as correlated processes, which has not yet been implemented in this literature. Secondly, each model is estimated separately for male and female children. Conditioning on gender will reduce biases due to discrimination in favour of males. Thirdly, each set of estimates takes account of family-specific, factors (e.g. health or socio-demographic factors of the parents) that are not observable to a researcher but that could be important in spacing or mortality equation. Finally, we choose the explanatory variables very carefully to reduce the problems posed by endogeneity bias (e.g., Rosenzweig and Wolpin, 2000).

The paper now considers the hypothesis in greater detail and describes the data and statistical model. The subsequent section presents and analyses the results. The paper is summarised in the final section.

2. HYPOTHESES, DATA AND METHODOLOGY

The Beckerian school of thought is essentially static in nature and does not take account of the sequential nature of childbirth. Two important extensions in this respect are the studies by Wolpin (1984) and Rosenzweig (1986). Wolpin 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. 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.

2.1. Hypotheses

Following Rosenzweig (1986), we determine child mortality as an indicator of child health outcome which depends on birth spacing, among other variables. It is argued that prior and posterior birth spacing reflect the age composition of consecutive siblings and is therefore an important aspect of competition among siblings for limited parental resources.[1]

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

[1] This works in conjunction with other possible channels, 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).

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 lies at the heart of these models and the reason for our interest in whether it is an empirically strong relationship.

Secondly, in societies characterised by pro-male bias, gender of the first-born, may also affect subsequent birth spacing and hence the health of subsequent children. This can be explained in terms of higher expected earnings of boys (Rosenzeweig and Schultz, 1982) and also the intrinsic randomness of a child's gender (assuming that parents cannot influence the gender of subsequent children). Thus parents characterised by son preferences are likely to strategically choose a shorter birth spacing if the first child is a girl, thus intensifying the competition among the siblings.

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.[2] 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.[3] 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

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[2] 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.

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

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)[4] in rural West Bengal has declined between 1981 and 1990, the state's own rate of decline in 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[5] 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 the birth spacing is less than 12 months but more than doubles when we consider non-first born children with the 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

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

[5] 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.

is a year or under, then subsequent females are over 30% more likely to die in the first 5 years.

2.3. Methodology

The unit of observation is a woman together with the birth history of all her children. The primary variable of our interest is child mortality, which among other things, depends on spacing between consecutive children. We can take account of both prior and posterior birth spacing although we do not observe prior birth spacing for first born children and posterior birth spacing for youngest children. We therefore concentrate on middle order children. We model each birth spacing as a hazard equation and child mortality is modelled as a probit equation. The following subsections discuss these equations.

Birth spacing

Posterior spacing: The log hazard rate of spacing from the time of birth of child *j* till the arrival of the next sibling (*NEXT*) is a function of calendar time (*T*(*t*)) and household (X_1) and individual child-specific (X_{2j}) characteristics and a family-specific[6] heterogeneity component *e* common to all children in a particular family. It is:

$$\text{Ln } h_j^{NEXT}(t, e) = \alpha_0 + \alpha_1 X_1(t) + \alpha_2 X_{2j}(t) + \alpha_3 T(t) + e + u_{Nj}$$

The subscript for the individual woman is suppressed for notational convenience. 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 :

$$\text{Ln } h_j^{PREV(}t, h) = \gamma_0 + \gamma_1 X_1(t) + \gamma_2 X_{2j}(t) + \gamma_3 T(t) + h + u_{Pj}$$

6 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.

The baseline hazard for each birth spacing equation is defined as piece wise 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 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).

Each birth spacing hazard equation depends on both parental and household characteristics. In particular, we include mother's age at first birth, mother's literacy, and a composite indicator of household assets.[7] We also include the characteristics of the children already born. [8] The choice and use of 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 dummy variable for whether the individual is Muslim because 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. We have obtained the first principal component of a number of asset variables and use this as a measure of wealth. Household assets as well as religion proxy household wealth in the Indian context. We also argue in our analytical model that the gender composition of the existing children (e.g. whether the first child is female) could significantly affect parental birth spacing decisions.[9]

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

[8] See Appendix 2 for the definition of these variables.

[9] Please note that when we jointly determine both prior and posterior birth spacing (along with child mortality) we include prior birth spacing as an explanatory variable for posterior birth spacing and vice versa.

Child mortality equation

We explain the probability of death using a mortality equation showing the probability that a child dies within 5 years of birth. The propensity to die is given by:

$$D_j^* = b_1X_{3j} + b_2PREV + b_3NEXT + d + u_{Mj}$$

The child dies if $D_j^*>0$ and death is recorded by the dummy variable, D_j, that takes the value 1 if the child has died. The vector of covariates X_{3j} includes exogenous variables. *PREV* and *NEXT* are the 'birth spacing' variables showing, respectively, the lengths of time since the births of the previous child and the next child. We adopt a probit specification that enables model family-specific differences, d, to be modelled as random effects. U_{Mj} is a random error.

We assume that the household chooses the number and age composition 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. The resulting problems of endogeneity in the mortality equation have yet to be solved. We have attempted to resolve the issue elsewhere using instrumental variables[10] but the use of weakly correlated instruments may exacerbate the problem. Although, like others, we do not fully solve this problem in this paper, we seek to reduce its impact by a number of devices that have not previously been implemented in this context.

10 Makepeace and Pal (2001).

We assume initially that the household is risk neutral and uses the expected value of the error term in its decision making.[11] Since the error term has zero mean, the optimal values of previous and subsequent birth spacing do not depend on the value of the error so we can treat these variables as uncorrelated with the error term. This would imply that the family made plans at the outset of its existence that it followed thereafter. However, plans are certainly revised over time so that birth spacing decisions are *de facto* taken at different points in time with different information sets. Nonetheless, any genuinely random component of the error term is uncorrelated with previous birth spacing because the value of this error is not known when the decision to have the child is made. However the value of the error term is likely to be correlated with subsequent birth spacing since decisions about the next child will be reviewed in the period after a birth using current information. This suggests that endogeneity will be more of a problem for the subsequent birth spacing. One solution is to accept that we cannot estimate its impact and omit it from the estimating equation. This would lead to problems if the lengths of the two birth intervals were correlated. However, our estimates of the impact of previous birth spacing do not appear sensitive to the omission of subsequent birth spacing. We would therefore argue that an increase in the time since the birth of the last child reduces child mortality is evidence of the importance of sibling rivalry.

The problem with this argument is that the random error affecting the life chances of a child when they are born is not known when the child is planned. Our specification assumes that family-specific factors are part of the error. Some of these, such as certain health problems with the child or family, may be unknown but others, such as attitudes to parenting, are surely known. Although we take the underlying heterogeneity into account, this is through a random effects specification.

[11] We could make the weaker assumption that the moments of the error term are independent of *PREV* and *NEXT*.

We have selected the variables in our model to take account of possible simultaneity problems. For example, inclusion of the number of siblings or number of sisters (as in Garg and Morduch, 1998) or even 'any sisters' in the child health function as in Butcher and Case (1994) may bias estimates because of endogeneity problems. This is because, number of sisters (or brothers) depends on the choice of family size. 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 and can therefore be used as an exogenous variable in the child health function. Similarly, we can use whether the first child died to indicate any serious family problems. Having a twin birth is positively correlated with having a large family. Rosenzweig and Schultz (1980) addressed this issue by dividing the number of twin births by the total number of pregnancies. Still there may be problems because total number of pregnancies is a choice variable. We consider if the current child is one of twins which in turn allows us to focus on aspects of competition among siblings for limited household resources. We also do not include the variable as to whether the child is the only one born to the parents since this is again chosen by the parents in the quantity-quality trade-off model.

In general, the probability that the *j*-th child dies will depend on a vector of characteristics, X_{3j}. Our model emphasised 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.

We include a number of variables reflecting various aspects of health. First, we include a dummy for whether the first child died. This may take account of 'death clustering'

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[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

such that families experiencing child death may have shorter birth intervals (Dasgupta, 1997) and therefore higher mortality rates. We have also included a variable to indicate if the current child is one of twins.[13] 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).

We hypothesized about the potential effects of sibling competition for limited parental resources on child health outcomes. In terms of the empirical mortality equation, these effects are captured by including a variable measuring the time to the birth of the next sibling or time since the birth of the previous child or both depending on whether the context child is the oldest, middle-order or youngest, on the grounds that rivalry may decline as the age gap increases. Thus parents can devote more time and effort to bring up a child if there is longer prior and posterior birth spacing, especially since this will also involve less maternal depletion.[14]

The provision of public services like water, sanitation, 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.[15] There may be gender differences so 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.

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causing problems of convergence. Hence, we could not include comparable characteristics of the father as we did for the mother.

[13] 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.

[14] We thus have a recursive system to determine the spacing hazard in terms of a series of exogenous variables and child mortality in terms of subsequent birth interval and other exogenous variables.

[15] 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.

Socio-cultural factors may also be important in parental allocation for investment in children. Religion may be considered to be an important factor in determining socio-cultural practices and to this end, we include controls for Muslim children. Finally, 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. Hence we proceed to consider separate male-female estimates of birth interval and child survival. We also indirectly consider the effect of birth order in child survival. This is done by considering separate estimates for (1) first-born and middle-order children and also (2) middle-order and youngest children.

Parent-specific unobserved heterogeneity

Since 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 (h,e,d) component and a child specific $(u_N,,u_P,u_M)$ component. 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, i.i.d. N(0,1) and independent of the family specific components. The correlation coefficients between the aggregate errors in the different equations are shown by r_{KL} where $K,L=N,P,M$.

The specifications of the birth spacing hazard (prior and posterior) and the mortality probit equations are summarised in Table 2. 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.[16] The conditional likelihoods are the probabilities of observed outcomes (the birth spacing hazard and child mortality equation for each child in the sample), conditional on the vector of unobserved heterogeneity components (h,e,d).

3. RESULTS

Tables 5 and 6 present, respectively, the estimates of the child mortality equation and the hazard equations for birth spacing for middle order children. The uncorrelated estimates of the child mortality equation show the results of estimating the probit independently of the birth spacing hazards. The correlated estimates allow for correlation between the unobservable variables in the two equations. The same factors are significant in both equations so the results are not qualitatively sensitive to the estimation technique. Moreover the magnitudes of most of the estimates are approximately the same in each set of results. Nonetheless the cross-correlations between the errors in the hazards and the mortality equation are significant so we concentrate on these 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 birth (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 the chances of survival.

3.1. Estimates of Child Mortality

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

Sibling age composition plays a central role in explaining child mortality. The main result is that an increase in the length of time either since the birth of the previous child or to the birth of the next child increases the chance of the child dying in the first 5 years of life. Secondly, being one of the twins increases the risks of mortality for both male and female children in our sample, again indicating the competition for limited resources both inside and outside the mother's womb. However, gender composition of older siblings, proxied by the gender of the first born, is insignificant in explaining mortality for both boys and girls.[17]

The role of education is confirmed since children of literate mothers are less likely to dye. The estimate for female children is 50% larger in magnitude so the beneficial effects of parental education are larger for daughters. Muslim girls have the same risk of dying as other girls but Muslim boys are less likely to die than other boys. Although its significance is less robust, an increase in wealth lowers the chances of a boy dying but has no significant impact for a girl. Death of the first child increases the mortality risks of subsequent female children, perhaps suggesting some pro-male bias in the response to this type of tragedy.

The two remaining variables, living in a rural district and age of mother at first birth, have no significant effect. Even if individuals living in rural areas are less well-off and have poorer access to government funded facilities, they do not fare worse than those in towns and cities. There is some evidence that age of mother at first birth does have an impact for first born children. When the model is estimated for first born and middle order children (excluding time since previous birth), an increase in the mother's age at first birth lowers the chance of a child of either sex dying. This may be because the mother's age variable acts a proxy for time since previous birth.

3.2. Estimates of Birth Spacing

[17] Though the variable may exert an indirect influence through its significant effects on posterior birth

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 the socio-economic variables, the hazard is lower if the mother’s age at first birth is higher and mother is literate. Also, the hazard is lower for more wealthy households, though the effect is significant for female children only. As predicted, the hazard is significantly higher for children born to Muslim parents, perhaps because of their general attitudes towards contraception use. 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 sample. Death of the first child however significantly shortens the prior spacing for both male and female children in our sample. Also, prior birth interval does not significantly affect posterior birth interval and vice versa for the posterior birth interval.

Finally, gender difference in birth spacing in our sample appears to be related to household wealth and gender of the first child.

3.3. Inferences

Thus these correlated estimates of birth spacing and child survival generally lend support to the central hypothesis of sibling rivalry in child survival in that shorter birth interval (prior and/or posterior) and twin births significantly enhance mortality risks among 0-5 year old male and female children. In general the parameter estimates from uncorrelated[18] and correlated models indicate similar pattern of results while the value of the log likelihood function is higher for the correlated estimates. In order to understand the extent of the bias in the uncorrelated estimates, we also compare the predicted probability of mortality for the middle order children,

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spacing, especially among male samples (see further discussion in section 3.3).

as summarised in Table 8. These predicted probability estimates not only suggest the 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 the mortality risks in our sample.

4. CONCLUDING COMMENTS

Within a sequential framework, the present paper examines the effects of sibling age composition on mortality risks of young boys and girls of different birth orders in the eastern Indian state of West Bengal. The innovations of the paper are as follows. First, we argue that competition among siblings for limited parental resources plays a significant role in child survival and is determined, among other things, by gender of the first child and age composition of consecutive children.

The empirical analysis based on the recent NFHS data from West Bengal employs a unique likelihood estimation technique to determine birth spacing hazard and mortality probit equations as correlated processes, allowing for mother/parents specific unobserved heterogeneity among first born, middle order and youngest male and female children. The explanatory variables are also carefully chosen to reduce the bias as far as practicable and include among various individual and household specific characteristics, the age (prior and posterior birth spacing) and sex (gender of the first child) composition of siblings to test the validity of competition among siblings on child mortality. These devices allow us to reduce the extent of endogeneity bias that has not previously been addressed 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

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

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. Mortality risks are worse for non-first born male children as the risk of having a subsequent sibling is higher if the first child is a female. Twin birth too significantly enhances the mortality risks of both male and female children.

Family resources have a direct effect both for determining birth interval and child survival such that wealth increases birth interval and thereby improves the likelihood of child survival. To the extent that the composite indicator of household assets, mother's age and literacy measures household economic status, it appears that children from better off families do better.

Household assets are also important determinants of gender differences in both birth spacing and child mortality in our sample. The other important factors explaining gender difference in child survival are region (rural) and religion, indicating the importance of socio-cultural practices in this respect.

These estimates suggest that there is 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 1. 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 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.nic.in/mohfw/popindi.html***

Table 2. Effects of sibling composition on child mortality

Numbers 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

Table 3. Specification of the Full Model

	Prior Birth Spacing	Posterior Birth Spacing	Child mortality
Intercept	√	√	√
Mother's age at first birth	√	√	√
Mother is literate	√	√	√
First child is a female	√	√	√
First child is dead	√	√	√
Twin birth			√
Prior spacing		√	√
Posterior spacing	√		√
Composite assets indicator	√	√	√
Muslim	√	√	√
Rural	√	√	√
Spacing 0-12 months	√	√	
Spacing 12-24 months	√	√	
Spacing >24 months	√	√	
Unobserved heterogeneity	*h*	*e*	*d*

Table 4: Sample characteristics - means and standard deviations of middle order children

	Male			Female	
Age of mother at first birth	23.16491	4.758555		23.22465	4.641633
Mother’s Literacy	0.290079	0.453873		0.297033	0.457026
Twin birth	0.020368	0.141279		0.017933	0.132729
First child female	0.494087	0.500047		0.515814	0.499831
First child died	0.268068	0.443026		0.257255	0.437192
Time since previous birth	30.5138	15.30044		30.59374	14.84545
Time to next birth	32.91754	17.05241		32.22237	17.52996
Princ. Comp. for Assets	-0.11564	0.89528		-0.12515	0.892179
Muslim	0.367608	0.482233		0.34431	0.47522
Rural	0.832457	0.373521		0.846104	0.360908
Number	3044			3067	

Table 5. Probit Estimates of Child Mortality: Middle-Order Children

	Uncorrelated estimate		Correlated estimate	
	Male	Female	Male	Female
Intercept	0.0058 (0.1824}	-0.139 (0.1887}	-0.135 (0.2701}	-0.5578** (0.2661}
Age of mother at first birth	-0.0034 (0.0065}	-0.0092 (0.007}	0.0052 (0.0119}	0.0027 (0.0125}
Mother's Literacy	-0.1636** (0.0773}	-0.2513*** (0.0796}	-0.1619** (0.0796}	-0.2409*** (0.0761}
Twin birth	1.2890*** (0.1889}	1.2519*** (0.1799}	1.3097*** (0.1937}	1.2554*** (0.1755}
First child female	-0.103 (0.0629}	-0.0431 (0.0653}	-0.1018 (0.0649}	-0.0305 (0.0624}
First child died	0.0913 (0.0686}	0.1558** (0.0738}	0.0984 (0.0708}	0.1488** (0.0699}
Time since previous birth	-0.0189*** (0.0024}	-0.0158*** (0.0023}	-0.0197*** (0.0034}	-0.0125*** (0.0031}
Time to next birth	-0.0117*** (0.0019}	-0.0108*** (0.0016}	-0.0123*** (0.0024}	-0.0084*** (0.0021}
Princ. Comp. for Assets	-0.0736* (0.04}	-0.0264 (0.045}	-0.0728* (0.0415}	-0.0288 (0.0427}
Muslim	-0.2287*** (0.0663}	-0.0383 (0.0696}	-0.2248*** (0.0688}	-0.0134 (0.0665}
Rural	0.0244 (0.0896}	0.0925 (0.0944}	0.0271 (0.0915}	0.0805 (0.089}
s_h			0.6230*** (0.0703}	0.5732*** (0.0657}
s_e			0.7019*** (0.0671}	0.6606*** (0.0609}
s_d	0.3426*** (0.0972}	0.4240*** (0.0709}	0.3619*** (0.099}	0.3326*** (0.0746}
r_{NP}			0.9998*** (0.0361}	0.9871*** (0.048}
r_{NM}			-0.1201* (0.067}	-0.1582* (0.0773}
r_{PM}			-0.0509** (0.0212}	-0.3136** (0.1201}

NOTE: Asymptotic standard errors in parentheses;
Significance: '*'=10%; '**'=5%; '***'=1%.

Table 6. Hazard Estimates of Birth Spacing: Middle-Order Children

	Posterior spacing Time to next birth			**Prior spacing** Time since previous birth	
	Male	Female		Male	Female
Hazard spline variables					
12 months and under	0.6605*** (0.1064)	0.7285** (0.1227)		0.6768*** (0.0935)	0.8038*** (0.0995)
24 months or less than 24 and more than 12 months	0.1652*** (0.0076)	0.1507*** (0.0073)		0.1504*** (0.007)	0.1471*** (0.0072)
More than 24 months	0.0099*** (0.0016)	0.0049*** (0.0018)		0.0145*** (0.0021)	0.0162*** (0.0023)
Intercept	-12.5562*** (1.2996)	-13.1185*** (1.4766)		-12.6850*** (1.1195)	-13.9177*** (1.1925)
Age of mother at first birth	-0.0151* (0.0091)	-0.0189* (0.009)9		-0.0091*** (0.001)	-0.0225** (0.0099)
Mother's Literacy	-0.0979 (0.065)	0.074 (0.064)1		0.0427 (0.0667)	0.013 (0.0626)
First child female	0.1018** (0.0501)	-0.0447 (0.05)		0.0514 (0.0529)	0.01 (0.0532)
First child died	0.0422 (0.0599)	-0.0479 (0.0592)		0.1580*** (0.0585)	0.1129* (0.0602)
Time since previous birth	0.0061 (0.0073)	0.0102 (0.0076)			
Time to next birth				0.0057 (0.0064)	0.0088 (0.0061)
Princ. Comp. for Assets	-0.0053 (0.0293)	-0.0670** (0.0315)		0.0292 (0.0332)	-0.007 (0.0345)
Muslim	0.0523* (0.0225)	0.0854** (0.034)		0.0568* (0.0249)	0.0980* (0.055)
Rural	-0.0433 (0.0695)	-0.0697 (0.07)		-0.1053 (0.0733)	-0.0336 (0.0713)

Table 7: 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 and posterior birth interval <=12 months		
Male	0.293555	0.310508
Female	0.240315	0.261792
If the child is a twin		
Male	0.549766	0.566160
Female	0.531657	0.549861

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
BORDER:	Birth order
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
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
MALE :	1 if the child is male and 0 otherwise