Infant Mortality and Child Nutrition in Bangladesh

Diane Dancer*, Anu Rammohan and Murray D Smith*****

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Abstract

The excess female infant mortality observed in South Asia has typically been attributed to gender discrimination in the intra-household allocation of food and medical care. However, studies on child nutrition find no evidence of gender differences. A natural explanation could be that in environments of high infant mortality of females, the surviving children are healthier, so that child nutrition cannot be studied independently of mortality. In this paper, we use data from the 2004 Bangladesh Demographic Health Survey to investigate if there are any gender differences in survival probabilities, and whether this leads to consequent differences in child nutrition. We argue the importance of establishing whether or not there exists a dependence relationship between the two random variables - infant mortality and child nutrition – and in order to detect this we employ a copula approach to model specification. The results suggest that male children have a significantly lower likelihood of surviving their first year. However, conditional on survival, they have better height-for-age Z-scores. From a policy perspective, household wealth and public health interventions such as vaccinations are found to be important predictors of better survival and nutritional outcomes.

JEL Classification: I12, C25, C51, O12

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* Discipline of Econometrics and Business Statistics, The University of Sydney, NSW 2006, Australia. E-mail: d.dancer@econ.usyd.edu.au

** Discipline of Economics, The University of Sydney, NSW 2006, Australia. E-mail: a.rammohan@econ.usyd.edu.au

*** Health Economics Research Unit, University of Aberdeen, UK, and Discipline of Econometrics and Business Statistics, The University of Sydney, Australia. E-mail: murray.smith@abdn.ac.uk

1. Introduction

1.1 Motivation and Context

Halving mortality rates, the prevalence of malnutrition amongst children and reducing gender disparities have been identified as key Millennium Development goals by the United Nations (UNDP, 2006). Bangladesh has made considerable progress towards these goals by improving survival probabilities for children and reducing malnutrition levels. For example, infant mortality rates in Bangladesh have improved from being 100 per 1000 children under 1 year of age in 1990 to 56 per 1000 in 2004. However, the nutritional status of children has shown a lesser degree of relative improvement with 48% of all children being described as 'underweight' in 2004, down from 66% in 1990, while 43% of children were 'stunted' in 2004, down from 65% in 1990 (UNDP, 2006).

Gender-induced disparities are of particular concern in Bangladesh given the well-documented evidence of the existence of a pro-son bias in South Asia, and its influence on child health outcomes (see Pitt *et al.*, 1990; Muhuri and Preston, 1991; Ahmad and Morduch, 1993; Morduch and Stern, 1997; Klasen and Wink, 2002). In an early study focusing on Bangladesh, Chen *et al.* (1981) found that girls residing in the Matlab region were less likely to be taken to medical practitioners, and were likely to have lower access to nutrients. Bairagi (1986) similarly found gender differentials in nutritional status with male children benefiting relatively more than female children from an improvement in economic circumstances in the Matlab region. Ten years later, Muhuri and Preston (1991) reported substantially higher female mortality among girls residing in Matlab if they had older sisters. They suggested that this points to selective neglect of individual female children. High mortality rates for female children have been attributed to the prevalence of ex-post gender discrimination against the girl child through discriminatory intrahousehold allocation of food, nutrition and medical care (see, for example, Bardhan, 1988; Behrman, 1988; Harriss, 1999).

If Muhuri and Preston's view persists and applies more widely to Bangladesh as a whole, then we should also observe poor nutritional status for female children relative to males. However, the link between both high child mortality and adverse nutritional status for females has not been well-established in empirical studies. A recent study of national-level survey data from a range of developing countries (including Bangladesh) finds little evidence of an anti-female bias in child nutrition measured using anthropometric measures (see Marcoux, 2002). This corroborates the findings of Ahmad and Morduch (1993) using the 1988 Household Expenditure Survey for Bangladesh, where they find limited support for gender bias in household expenditures, although the sex ratios in the sample data are dramatically skewed.

According to Morduch and Stern (1997) empirical support for sex bias is stronger in studies using aggregate data than those that use household-level behavioral analyses. These sentiments are echoed by Marcoux (2002) who notes that the belief in an anti-female bias in the intrahousehold allocation of food is a consequence of the extrapolation of results from studies of a small part of a country rather than using national data, as well as in disregarding the body of evidence of contrary results. Yet despite the evidence of excess infant mortality of female children on the one hand, and on the other hand a lack of empirical support for gender bias in recent studies on nutrition, little work has been carried out linking infant mortality to nutrition outcomes. A key problem has been a lack of availability of gender disaggregated nationally representative data on infant mortality and child nutrition. Children who survive are likely to differ systematically from those who do not and ignoring the selection bias is likely to result in biased and inconsistent estimates.

In this paper we use the nationally representative *Bangladesh Demographic Household Survey 2004* to investigate if there are any gender differences in survival probabilities, and whether this leads to consequent differences in child nutrition. Infant mortality is defined as mortality in the first year of life, and child nutrition is measured using two commonly used anthropometric measures: weight-for-height and height-for-age (see Waterlow *et al.* 1997). Malnutrition on account of a child having a low 'height-for-age' causes *stunting*, while a low 'weight-for-height' is associated with *wasting*. A child's 'height-for-age' is considered a stock measure, an indicator of their long run nutrition status reflecting the child's past nutritional experience. A child's 'weight-for-height' is regarded as a flow measure of short run or current nutritional status. Both measures are less sensitive to changes at the extremes of distributions of these variables and they facilitate comparisons across measures that exhibit different variability in terms of units of measurement.

This paper contributes to the existing research in two main ways. We show that, in an environment of high female child mortality and a preference for sons, child nutrition cannot be studied independent of mortality. In doing so, we provide a better way of identifying the determinants of infant mortality and child nutrition than studies that examine them independently. In econometric terms, our study provides a methodological improvement on previous research by allowing a preferred fit to emerge amongst a variety of self-selection models that differ according to copula specification. The copula method applied to our self-selection model describes the estimation method that applies to any set of assumptions that are imposed: these being assumptions on both the margins and the joint behaviour of the margins.

The copula method describes an econometric technique that has the advantage of allowing estimation of any self-selection model. Standard models of selectivity permit a limited number of statistical distributions. The Heckman model (Heckman, 1979), for example, imposes bivariate normality onto both margins, with each margin itself being normally distributed. The analysis in this paper effectively extends the range of available modelling options from a mere one-only (the Heckman) to include a range of other possibilities. This approach to studying the link between infant mortality and child nutrition is unique to this literature. For the dataset studied here we show that the Frank copula provides the superior fit, improving over the traditional Heckman selection model that corresponds to specification of a Normal copula. Further, the strong correlation estimated between infant mortality and child nutrition in all our preferred models supports our decision to adopt a sample selection approach.

The analysis in this paper uses a nationally representative sample of Bangladeshi children aged 0-60 months. We analyse the factors influencing infant mortality, and examine if there is any evidence of an anti-female bias. Bangladesh is an ideal context in which to research these issues because despite having a relatively high infant mortality rate and poor child nutritional outcomes for children, recent national-level data show a decline in both mortality rates for children and a lack of anti-female bias in the infant mortality rates. This is in contrast to the pattern observed in other South Asian countries such as India, where a strong male preference is thought to result in nutritional and medical neglect of female children.

The next section describes the dataset that is used in the analysis and in Section 3 we present the relevant descriptive statistics. Section 4 describes our econometric methodology and the results are presented in Section 5. Finally, our main conclusions are presented in Section 6.

2. Data

The data used in this analysis are drawn from the nationally representative 2004 *Bangladesh Demographic Health Survey* (BDHS) of 11,440 ever-married women aged 10-49 years and 4,297 men aged 15-54 years found in 10,500 households covering 361 sample points (clusters) throughout Bangladesh; 122 in urban areas and 239 in rural areas. This survey is the fourth in a series of national-level population and health surveys conducted as part of the global Demographic and Health Surveys (DHS) program. The survey consisted of two parts: a household-level survey of women and men and a community survey around the sample points from which the households were selected. Financial support for the BDHS survey was provided by the United States Agency for International Development (USAID)/Dhaka. It was implemented through a collaborative effort of NIPORT, Mitra and Associates, and ORC Macro.

The survey utilized a multi-stage cluster sample based on the 2001 Bangladesh Census and was designed to produce separate estimates for key indicators for each of the six divisions of the country—Barisal, Chittagong, Dhaka, Khulna, Rajshahi and Sylhet. The dataset is ideal for our analysis as it contains detailed information on household structure, labour market participation, asset ownership, health and educational characteristics for all the household members.

Our analysis is based on data from the Women's Survey component of the 2004 BHDS reported on 5172 children aged 0 - 60 months for whom complete information is available with regard to health, parental, household and community characteristics. We restrict our analysis to only those households that had children born in the 5 years prior to the survey. This restriction ensures that there is compatibility between the infant mortality figures and our child nutrition measures, which are only available for children under the ages of 60 months. Furthermore, this also reduces the potential for recall bias and measurement error.

Infant mortality is defined here as the death of a child before the age of one year. From a total sample size of 5172 children that were born in the five years prior to the survey, 461 did not survive, and 4711 were alive and at least one year of age at the date of the survey. As in Arulampalam and Bhalotra (2006), we exclude from our analysis children who were less than 12 months of age, but alive at the time of the survey since these children have not had a full year's exposure. We further note that there is no data on anthropometric measures for children who have died and we are unable to determine the cause of death. Since our infant mortality measure is defined as the death of children under the age of 12 months, we exclude children who may have died between the ages of 12-60 months.

We measure child nutrition using two anthropometric measures: a child's height-for-age and weight-for-height, both are expressed in standard deviations (Z-scores) from the mean of the reference population, this being the commonly used US National Center for Health Statistics (NCHS) standard as recommended for use by the World Health Organisation (WHO). The height-for-age Z-score measures the child's height according to age, this being an indicator that reflects the cumulative effects of growth deficiency and so is designed to measure long-term nutrition. The weight-for-height Z-score measures the child's weight according to height, where this indicator has been used to monitor the growth of children and is typically regarded as a measure of short term rather than long term health status. Both anthropometric measures are influenced by a number of factors including chronic insufficient food intake, frequent infections, sustained incorrect feeding practices and low socio-economic family status. However, since data on food intake is unreliable, these anthropometric measures are widely regarded by nutritionists as a reliable indicator of malnutrition.

3. Descriptive statistics

Table 1 presents descriptive statistics for the main variables of interest for our sample of 5172 children, disaggregated by whether or not the child is alive and then by gender. In our dataset there are 4711 children who were alive and at least one year of age at the time of the survey, with boys numbering a slender majority. With regards to the anthropometric measures, we observe that children in the sample have mean height-for-age Z-scores slightly less than 2 standard deviations below the NCHS mean, whereas the weight-for-age Z-scores are slightly larger than 1 standard deviation below the NCHS mean. Hence, stunting appears to be a bigger problem in Bangladesh, relative to wasting. The sample is predominantly rural with only 30% of the children living in urban areas.

A non-parametric kernel density plot of 'height-for-age' and 'weight-for-height' for all those in the 'alive' group appears in Figures 1 and 2 respectively. The estimated densities have a Gaussian-like appearance, providing justification for imposing that assumption on the margins in the self-selection models that are to follow.

The rest of this section describes the explanatory variables used in the analysis. To compare across the different models, we include the same set of explanatory variables across all our models. Child survival and child nutritional outcomes both depend on a set of child-specific, household demographic, economic and parental characteristics and a set of health inputs. The child-specific variables included among our regressors are the child's gender, birth-order, and an indicator variable for whether or not the child is part of a twin. For birth-order we use the absolute birth order of each child in the household to compute five dichotomous birth order dummy variables - second-born, third-born, fourth-born, fifth- to tenth-born (with the first born child being the base).

In a recent paper using Indian DHS data, Arulampalam and Bhalotra (2006) introduce the idea of 'scarring'. They argue that a child's death 'scars' the family, so that parents become more vulnerable. They use the previous sibling's survival status to capture the effects of scarring. We also control for this effect in our estimation by including an explanatory variable on whether the previous pregnancy resulted in a miscarriage, still-birth or death. There are 507 such mothers in our dataset.

A shortcoming of this dataset is that it contains no information on wages, household expenditure patterns and community characteristics. However, the dataset contains a wealth index which divides households into 5 wealth quintiles, which is generated using the principal components analysis. It is a composite measure of the cumulative living standard of a household, which places individual households on a continuous scale of relative wealth. The wealth index is divided into population quintiles, with the lowest quintile representing the poorest 20 percent and the highest quintile representing the wealthiest 20 percent households (see Filmer and Pritchett, 2001). These wealth quintiles have the advantage of providing a reasonably reliable measure of the household's economic status, and it is not affected by the endogeneity and transitory nature of labour income.

We include a set of dummy variables for the education levels of the mother and father. Since father's education levels are generally higher, we include an additional category 'higher secondary and above' for fathers. There is a plethora of research linking improvements in maternal education to better child health outcomes (see Glewwe, 2000; Behrman, 1988; Strauss and Thomas, 1995; Strauss, 1990; Thomas *et al.*, 1991). Improvements in maternal education levels improve child health outcomes through several channels including: safer health and hygiene practices, more exposure to information and a better ability to act on the information (Alderman *et al.* 2006).

From our descriptive statistics (Table 1), it is noteworthy that a larger proportion of girls from the poorest wealth quintile (wealth 1) have died relative to boys (26%). A comparison of male and female children who are alive at 12 months also shows that girls have poorer height-for-age Z-scores relative to boys.

In the group of children that are alive, we note that the proportion of children affected by a previous death or miscarriage in the household is roughly similar across the two genders. We further observe that boys are more likely to be born higher in the birth order.

The household demographic and regional characteristics are roughly similar across male and female children. A comparison of the household characteristics of the children who survived to those that died shows that the latter group on average lived in poorer households, had a greater proportion of mother's who are working, and had a greater percentage of mothers who are illiterate (48%) when compared to children in the 'alive' sample (37%). Similarly, we see that nearly 47% of fathers of children that have died during their first year of life have no education.

Since only 18% of the mothers in the sample are employed, we include an indicator variable for whether or not the mother is working. Empirical research on the influence of maternal work on child nutrition is ambiguous (see Leslie, 1988 and Glick and Sahn, 1995). However, given the traditional role of women in Bangladeshi society, it is likely that working women come from less affluent households. For father's employment status we include a set of 3 dummy variables

We observe low levels of maternal education levels in the sample, with 37% of the mothers reporting no educational achievement whatsoever. There are 39% of fathers who report no educational achievement, but employment amongst fathers seems universal with all reporting being in employment. Nearly 45% of the fathers are employed as manual labourers with agricultural workers constituting 25% of the sample

In our dataset maternal socio-economic status is measured using a series of self-reported qualitative questions measuring social and economic autonomy, such as the mother's control over economic resources, her knowledge and ability to make decisions on health care both with regards to herself and her children, and her freedom of movement. These issues are important in Bangladesh as women have traditionally had very little decision making power in the household, and also face several restrictions on working outside, travelling alone without being accompanied by male or elderly relatives, and in accessing health care provided by males. It is interesting to note that mothers generally have low levels of autonomy with only 47 % having the knowledge of where to locate medical facilities for treatment for themselves. However, while only 15% of the mothers know where to find medical help if their child is ill, we observe that only 10% of the mothers need permission to be allowed to access health care.

Immunization of children has long been established as a cost-effective way to improve child health. Unlike other policy interventions, however, immunization may also be regarded as a measure of personal illness control, and could be indicative of parental motivation to ensure a child's well-being and health (see Pande, 2003). The data on vaccinations are only available for children that are currently alive, hence we include an indicator variable for whether or not the child has been vaccinated in our child nutrition model.

4. Econometric Methodology

In this section, we set out our econometric model that takes into account the selection issues relating to child nutrition. In particular, we examine the probability of the child surviving, and contingent upon the probability of survival, we then estimate the factors affecting child nutrition for the sample of children aged between 12-60 months.

In this paper the 'copula approach' to model specification is utilised in the context of binary models designed to allow for possible data selectivity. Selectivity is frequently a problem with microeconometric data whereby underlying individual circumstances can themselves influence the observations collected on random variables. Statistical models of increasing complexity have been constructed to account for selectivity in its various guises, should it be present, and a number of these are discussed in texts such as Amemiya (1985), Maddala (1994) and Lee (1996). The copula approach is a modelling strategy whereby a joint distribution is induced by specifying marginal distributions, and a function that binds them together: the copula. The copula parameterises the dependence structure of the random variables, thereby capturing all of the joint behaviour. This then frees the location and scale structures to be parameterised through the margins, one at a time. Most importantly, the copula approach permits specifications other than multivariate normality, although it does retain that distribution as a special case.

Of particular interest here is the self-selection model in which a binary indicator S governs whether or not an observation is generated on a second random variable Y. Selectivity arises if S and Y are correlated. Importantly, our main concern is whether child nutrition can be studied independently of infant mortality in environments with high mortality rates. *A priori* it is difficult to predict whether there will be a positive or negative dependence between survival and child nutrition. For example, we expect a positive dependence between survival and a child's relative nutritional status if surviving children are inherently stronger. However, if greater resources are being directed to ensure child survival, and there is relative neglect of the children that have survived, then it is possible to conceive of a situation in which survival probabilities are negatively correlated to child nutrition.

Following the general copula modelling procedure described in Smith (2003), we embed the self-selection model within a latent utilitarian framework that will later be related to the observed variables through transformation rules. Let

$$Y_1^* = \text{probability of a child surviving to 12 months of age} \sim N(x_1'\beta_1, 1) \qquad (1)$$

which can be written as a linear function of variables that affect the probability of a child survival and

$$Y_2^* = \text{a child's relative nutritional status} \sim N(x_2'\beta_2, \sigma^2) \qquad (2)$$

where evidence for the Normality assumption for Y_2^* appears in Figures 1 and 2 for both height-for-age and weight-for-height analyses. Note that x_i is a $(k_i \times 1)$ vector of covariates of Y_1^*, and vector β_i $(k_i \times 1)$ and scalar σ are unknown parameters to be estimated (i = 1, 2). The joint distribution function of the latent variables $Y_1^* = y_1^*$ and $Y_2^* = y_2^*$, is expressed using Sklar's representation, namely,

$$F_{12}(y_1^*, y_2^*) = \Pr(Y_1^* \le y_1^*, Y_2^* \le y_2^*) \qquad (3)$$
$$= C_\theta \left(\Phi(y_1^* - x_1'\beta_1), \Phi\left(\frac{y_2^* - x_2'\beta_2}{\sigma} \right) \right)$$

where C_θ denotes the copula function that binds together the two margins, and Φ denotes the distribution function of a standardised Normal random variable. Having specified the margins (1) and (2), the final step of the copula approach is to specify a copula function. To illustrate, two contending specifications worthy of mention here for their relevance to the empirical results presented below are the Gaussian copula and Frank's copula. The former is given by

$$C_\theta(u,v) = \Phi_2\left(\Phi^{-1}(u), \Phi^{-1}(v); \theta\right) \qquad (4)$$

where Φ_2 denotes the distribution function of a standardised bivariate Normal random variable with Pearson correlation coefficient θ, the dependence parameter. Note that $(u,v) \in \Re^2$. Using (1) and (2) in (4) finds the joint distribution function of $Y_1^* = y_1^*$ and $Y_2^* = y_2^*$ given by

$$\Phi_2\left(y_1^* - x_1'\beta_1, \frac{y_2^* - x_2'\beta_2}{\sigma}; \theta \right) \qquad (5)$$

leading to the so-called Heckman self-selection model (see Amemiya, 1985) that is pre-programmed in a number of popular econometric estimation packages.

Frank's copula, this time defined for $(u,v) \in [0,1]^2$ and dependence parameter θ now such that $-\infty < \theta < \infty$, is given by

$$C_\theta(u,v) = -\theta^{-1} \log\left(1 + \frac{(e^{-\theta u} - 1)(e^{-\theta v} - 1)}{e^{-\theta} - 1} \right) \qquad (6)$$

For this copula, negative/positive values for θ imply a negative/positive dependence relationship between Y_1^* and Y_2^*; independence occurs if $\theta \to 0$.

The transformation rules that relate (Y_1^*, Y_2^*) to the observed variables (S,Y) in the self-selection model are

$$S = 1\{Y_1^* > 0\} \text{ and } Y = 1\{Y_1^* > 0\} Y_2^* \tag{7}$$

where $1\{A\}$ denotes the indicator function, taking value 1 if event A holds, and 0 otherwise.

In effect, nutritional status Y_2^* can be observed only when $Y_1^* > 0$. Clearly, a child's survival is represented by the Bernoulli variable S, and this random variable derives its properties from those of Y_1^*. Note that when $S = 0$; (i.e. death occurring at an age less than one year), Y_2^* cannot be observed, and Y is assigned a dummy value of 0.

Using data on n children (s_j,y_j), j=1,…,n, with outcomes assumed to be mutually independent, results in the likelihood function

$$L = \prod_{s=0} \Phi(-x_1'\beta_1) \prod_{s=1} \left(\sigma^{-1}\phi\left(\frac{y - x_2'\beta_2}{\sigma}\right) - \frac{\partial}{\partial y} F_{12}(0, y) \right) \tag{8}$$

where, for convenience, the index j has been dropped and ϕ denotes the density function of a standard Normal random variable. Smith (2003) lists the likelihood functions corresponding to the Gaussian and Frank copulas, amongst others.

The covariates x_1 and x_2 are fixed across all fitted models; there being 36 variables in x_1 and 45 in x_2. In all tables, estimates of β_1 for the x_1 set are followed by estimates of β_2 for the x_2 set.

A variety of self-selection models were fitted using differing copulas including both the Heckman and the Frank models; however, the only ones reported here are the Heckman and the Frank models. Results for the unreported models may be obtained from the authors upon request. The Heckman is the well-known, pre-packaged model available in many econometric computing packages, but it has the disadvantage of imposing bivariate Normality. For our data, the model fitted using Frank's copula was better-performed across all but one model.

5. Results

The results of our analysis for the preferred models are presented in Tables 2 and 3.[1] All the estimations were carried out for the full sample that combined together boys and girls, as well as disaggregated by gender to allow for gender-specific comparisons. In all cases except the weight-for-height estimation for boys, the Frank model provides the better fit amongst a range of models all of which differed according to the copula function used. Table 2 presents

[1] The results for other selection model are not presented here but are available from the authors upon request.

the estimates for the Frank model for height-for-age; we present estimates for the probability of the child being alive, for the full-sample, female-only and male-only samples. Table 2 also reports the second-stage estimates of height-for-age for the sample of children that are alive and exceeding 12 months of age at the time of the survey, by full sample and disaggregated by gender. In Table 3, results are presented for the weight-for-height model, for which two of the three fitted models report estimates obtained using the Frank copula while the Heckman model constructed from the Gaussian copula is presented for the male-only sub-sample.

A comparison of the maximised log-likelihoods for the full-sample between the Heckman (-8705.07) and Frank models (-8693.24) prefers the latter over the former in the height-for-age estimation, with this simple method permitted as a valid criterion (equivalent amongst a range of information criteria) for selecting between these two non-nested models because both contain the same number of parameters. The Frank and the Heckman nest the Independence model (maximised log-likelihood -8706.28), but whereas the Frank model shows significant improvement (e.g. the Wald Test on the significance of Kendall's $\tau = 0$ yields a t-statistic of -8.28), the same statistic for the Heckman is not as strong (the Wald Test on the Kendall's $\tau = 0$ yields a t-statistic of -1.95).

In the weight-for-height estimations, we similarly observe that use of the Frank copula provides a better fit in the full model and in the female-only sample. For example, in the Frank model, the Wald test on Kendall's $\tau = 0$ yields a highly significant t-statistic of -10.07, the same statistic for the Heckman model yields an insignificant t-statistic of -0.97 on a Wald Test on Kendall's $\tau = 0$). We note that in general although the coefficients have similar signs in the Frank and Heckman models, the size of the coefficients are larger in the Frank model. To keep the discussion tractable we first discuss the probability of survival and then the results from the two nutrition regressions.

Probability of Surviving

Our estimation results show that, in all our models, female children are significantly more likely to survive. This result is surprising, as it is contrary to previous studies from India and Bangladesh that find evidence of excess female infant mortality rates (see Maitra *et a*l., 2006 for India; Muhuri and Preston, 1991 for Bangladesh). Our findings however, are in keeping with the raw statistics from the BDHS 2004, and with the biological trends in infant mortality, where male children have a lower propensity to survive relative to females. We further note that unlike previous studies from Bangladesh, such as Muhuri and Preston (1991) that is restricted to the Matlab region of Bangladesh, the data used in our analysis is nationally representative.

With regard to child specific characteristics, we see that, in all our models, regardless of the child's gender, being part of a twin significantly reduces the probability of survival. Interestingly, being born later in the birth-order appears to be beneficial in general with regard to child survival probabilities. We observe here that, relative to the first-born child, a child who is born second, third or fourth in order has a significantly higher likelihood of survival. This would suggest that, with each additional child, parents are getting better informed on ways to improve survival probabilities.

To get a better perspective of the link between birth-order and gender we estimated included into the full sample height-for-age and weight-for-height models interactions of the gender variable with each of our birth-order variables. In this instance (estimation results are not tabulated here but are available upon request), both models indicate that a male child who is born fifth or higher in the birth order is less likely to survive relative to a first born female child. However, the interaction terms jointly have no significant influence on the height-for-age Z-score, while the Z-score for the weight-for-height model experiences only a small positive effect for one interaction term: male children born fifth or higher in birth-order.

In respect of the scarring issue raised by Arulampalam and Bhalotra (2006), our analysis shows that the death of a previous sibling (due to still-birth, miscarriage or infant death) has no significant effect on an infants likelihood of survival.

From the set of variables that capture the household demographics, we observe that an increase in the number of children under five years of age has a significantly positive effect on survival across all our models. This is consistent with the better survival probabilities observed for later-born children.

The explanatory variables have a differential impact on the male and female samples. Male children from larger households are more likely to survive in both the height-for-age and weight-for-height models, with no effect on the female sample. Further, parental education has no influence on the likelihood of survival of male children. However, we see that, in all our models, relative to having a mother with no education, children (in the full sample and in the female sample) whose mother's have at least a primary or a secondary education are significantly more likely to survive. We similarly see positive effects from father's education for both these two groups. In particular, having fathers with higher than secondary level of education significantly increases the likelihood of survival of children in the full sample and females. Furthermore, an increase in the proportion of daughters in the household has a significant negative effect on the survival probability of male children. For female children however, we see the opposite effect. This result is in keeping with Garg and Morduch's (1998) finding for Ghana where having female siblings had a beneficial impact on the health of girls.

Nutrition measures

With the exception of the male sample (for weight-for-height) in the Heckman model, in all our other preferred models (all of which are based on using the Frank copula) there is a negative dependence estimated between child survival and nutrition measures; see the estimates reported for Kendall's τ measure of association. One possible explanation for this result, even given the negative coefficient observed on the gender dummy for survival in the full sample, may be that parents may be directing their resources to ensuring that their children survive, but perhaps neglecting the nutrition of those children that have survived. This would be consistent with the raw data from Bangladesh which shows a dramatic decline in infant mortality over the last few years, with little change in child nutrition. For male children in the Heckman model of weight-for-height, we observe a significantly positive Kendall τ, indicating that male children are positively selected into the sample.

Importantly, our results indicate that although the male child has a significantly lower probability of surviving, conditional on having survived past their first birthday, a male child has a statistically significant likelihood of having a better height-for-age Z-score (Male 0.18 standard deviations better than that of a female child), but that same gender effect in the weight-for-height Z-score model is not statistically significant (Male 0.01). It is possible that given the higher likelihood of male deaths in infancy, parents are compensating male children by providing them with greater access to nutrition and medical care, at least in the longer term.

We observe significantly negative birth-order effects in the height-for-age models, with later born children having poorer height-for-age Z-scores relative to earlier born children. This is in contrast to the better survival probability of later born children. This effect is particularly strong for a child born fifth or higher in the birth-order. As Table 2 shows the height-for-age Z-score of a child born fifth or higher in the birth-order is nearly 0.19 standard deviations below the first born child in the full sample, with similar magnitudes in the female-only and male-only samples. However, in the weight-for-height models we fit, the sign of the coefficients differ across genders. In the female-only sample, relative to a first-born child, a female child born second, third or fourth in the birth-order has a significantly lower weight-for-height Z-score, which worsens monotonically with each subsequent birth-order. With the male child however, we see the opposite effect. Our estimation results show that a male child who is born second or third in the birth order has a weight-for-height Z-score that is over 0.11 standard deviations above a first-born male, with a male child born fifth or higher in the birth-order having a Z-score that is 0.15 standard deviations above a first-born male child. Similarly, while an increase in the number of children under 5 years of age has a significantly negative effect in the sample of girls, we see that this variable has a positive effect on male children. As weight-for-height is a measure of short-term nutrition and is responsive to short-term changes in diet, this suggests that, despite resource constraints, young male children receive good nutrition.

While the death of a previous child is not influential in affecting survival probabilities, we note that in the sample of children who survived, the weight-for-height Z-scores are nearly 0.10 (0.20) standard deviations higher in the full (female) samples. One possibility for this is that there may be a learning effect with the death of a previous child, with parents getting better information on improving nutrition and reducing the incidence of disease. Therefore, unlike Arulampalam and Bhalotra's (2006) study, we find positive effects due to scarring. We note, however, that the evidence from our fitted models is not particularly strong, for this variable has no effect on male children and only affects the female-only sample in height-for age, where the effects are positive and significant at merely the 5% level.

Household economic characteristics have mixed effects on child anthropometrics. While father's occupation has no significant effect on height-for-age in any of our models, household wealth emerges as being highly significant in improving height-for-age Z-scores. For example, relative to the base category (wealth quintile 1 - the poorest), belonging to each higher wealth quintile significantly improves a child's height-for-age Z-score, irrespective of whether we look at the full-sample estimates or the gender-disaggregated estimates. In particular, a child born into the wealthiest quintile has a height-for-age Z-score that is 0.63 standard deviations higher than that of a child from the poorest wealth quintile. We note that the sizes of these wealth effects are larger for the male-only sample. For example, a male child born in the wealthiest quintile has a height-for-age Z-score that is 0.69 standard deviations higher than that of a male child from the poorest wealth quintile. However, for female children, belonging to the highest wealth quintile has a slightly smaller effect on height (0.58 standard deviations). Since height-for-age is a stock measure reflecting long-term nutrition, this result would indicate that poverty has a detrimental effect on a child's long-term nutrition.

Surprisingly, in terms of shorter-term nutrition outcomes reflected in our weight-for-height estimates, the wealth variable displays far less importance. It is only significant for children in the third quintile in both the full-sample model and the male-only sample model, the remaining coefficient estimates are all individually insignificant from zero.

Our estimations show that while father's occupation has no influence on the male-only sample, female children whose fathers are employed as office workers or manual labourers have weight-for-height Z-scores that are approximately 0.10 standard deviations higher than those of girls whose fathers are employed as agricultural workers.

In the set of variables that capture mother's autonomy and health characteristics, we observe that mother's weight has a large, positive and significant effect on the weight-for-height Z-score of children across all our preferred models, whereas mother's height only has significantly positive effects on the height-for-age Z-scores of children in the full-sample and for males-only.

It is interesting to note that mother's autonomy variables have greater impacts in the full sample and the female-only model, with only one instance occurring in the male-only model when one of these variables is significant. There is a persistent positive and significant effect across all weight-for-height Z-score models in regard of the variable representing a mother who knows where to get help if their child is ill. However, if the mother needs permission to spend money or permission to access health care, it reduces the female child's weight-for-height Z-score by 0.14 and 0.11 standard deviations, respectively.

Finally, in a recent study from Bangladesh, Bishai *et al* (2003) have shown that public health interventions in the form of vaccination against infectious diseases are a key, low-cost health intervention to improve child heath. Our results also confirm the positive effects of immunization. Across all our samples, children receiving vaccinations (up to 4 can be administered) experience significantly better height-for-age outcomes per vaccination; boys especially benefiting almost two-fold over girls. Given that height-for-age is a stock measure of long-term health, this is a cost effective public health intervention to improve child nutrition. However, we note that our modelling suggests no improvement in a child's short-term nutrition as a result of undergoing vaccination, this being evidenced by insignificant estimates in each of our weight-for-height models.

6. Conclusions

In this paper we have examined the link between infant mortality and child nutrition in Bangladesh. We have argued that, in an environment of high infant mortality, child nutrition cannot be studied independently of infant mortality. For this reason we estimated a sample selection model, the self-selection model, using a copula approach to modeling. Our analysis confirms that dependencies are important; implying that to study child nutrition independently of infant mortality will lead to inconsistent estimates. From an econometric standpoint, we have shown the importance of casting more widely for potentially better-fitting models than the ubiquitous Heckman model.

Our results may seem somewhat surprising in the context of the well-documented excess female mortality rates in South Asia. However, the negative dependence between survival probability and nutritional outcomes of the children that have survived painted a different picture. Here we found that, relative to male children, female children were more likely to have poor weight-for-height and height-for-age Z-scores. Hence, the higher infant mortality rates for male children may simply be a manifestation of their lower survival probabilities biologically. Other factors such as the availability of sex-selective abortion technology, and the recent dramatic fall in infant mortality and fertility rates in Bangladesh may also be contributing to the better survival probability of female children.

Our analysis finds no evidence of any gender discrimination in the survival probability of children under the age of one year. However, when we examined the anthropometric measures of children who were aged at least one year at the time of the survey, we saw that male children had better height-for-age Z-scores relative to female children. Mother's, but not father's education, was shown to be an important predictor of child survival, particularly for female children. Economic variables such as household wealth were shown to have an important positive effect on height-for-weight.

From a policy perspective our results showed that public health interventions in the form of vaccinations are influential in improving a child's longer-term health.

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Table 1- Descriptive statistics

	Children alive at 12 months N=4711		Children who died N=461	
Sample Size	**2390**	**2321**	**249**	**212**
Child characteristics	**Male**	**Female**	**Male**	**Female**
Height-for-age	-1.905 (1.29)	-1.975 (1.30)		
Weight-for-height	-1.126 (0.90)	-1.131 (0.88)		
Age – in months	35.428 (13.71)	34.556 (13.85)		
Twin			0.100	0.094
Number of vaccinations	3.708 (0.78)	3.610 (0.93)		
Birth Order 1	0.332	0.329	0.426	0.382
Birth Order 2	0.250	0.256	0.193	0.203
Birth Order 3	0.181	0.188	0.129	0.137
Birth Order 4	0.106	0.112	0.100	0.099
Birth Order 5 and higher	0.131	0.115	0.153	0.179
Previous child died or mother miscarried	0.103	0.103	0.092	0.104
Household demographic characteristics				
Muslim	0.913	0.911	0.936	0.943
Household size	6.451 (2.92)	6.531 (3.09)	5.743 (2.77)	5.358 (2.57)
Number of children under 5	1.521 (0.74)	1.517 (0.75)	0.956 (0.79)	0.953 (0.80)
Mother's age yrs	26.215 (6.30)	26.236 (6.16)	26.137 (7.30)	26.420 (6.90)
Mother's weight kgs	44.814 (7.66)	45.104 (7.89)	44.150 (8.53)	43.811 (7.39)
Mother's height cms	150.443 (5.51)	150.435 (5.41)	149.534 (5.49)	150.405 (5.56)
Mother's education – primary	0.321	0.316	0.293	0.274
Mother's education – secondary	0.308	0.317	0.265	0.217
Mother not working	0.815	0.816	0.839	0.731
Father's education – primary	0.273	0.275	0.241	0.269
Father's education – secondary	0.233	0.241	0.249	0.208
Father's education – higher than secondary	0.094	0.110	0.064	0.042
Household economic characteristics				
Father's occupation – office worker	0.285	0.320	0.285	0.236
Father's occupation – manual labourer	0.455	0.446	0.466	0.514
Wealth quintile – 1st	0.243	0.223	0.261	0.335
Wealth quintile – 2nd	0.203	0.185	0.157	0.189
Wealth quintile – 3rd	0.182	0.188	0.229	0.184
Wealth quintile – 4th	0.186	0.171	0.165	0.108
Wealth quintile – 5th	0.187	0.233	0.189	0.184
Access to piped water for drinking			0.084	0.075
Mother's autonomy				
Where to go to access health care for self	0.464	0.468		
Needs permission to spend money	0.138	0.135		
Knows where to get help if child is ill	0.143	0.166		
Needs permission to access health care	0.065	0.065		
Needs permission to spend money on health	0.094	0.103		
Regional characteristics				
Urban	0.295	0.307	0.297	0.283
Chittagong	0.216	0.209	0.249	0.160
Barisal	0.113	0.112	0.100	0.085
Khulna	0.128	0.130	0.072	0.108
Rajshahi	0.204	0.190	0.165	0.189
Sylhet	0.124	0.132	0.181	0.175

Notes: Standard deviations for the continuous variables and the count variables appear in parentheses. Only proportions are shown for binary 0-1 dummy variables.

Table 2: Maximum Likelihood Estimates: Height-for-Age

	Stage 1- child alive			Stage 2- height-for-age estimates		
	Full sample	Female	Male	Full sample	Female	Male
Child characteristics						
Male	-0.139**			0.180**		
Twin	-1.562**	-1.619**	-1.660**			
Birth Order 2	0.213**	0.361**	0.441**	0.064	0.038	0.090
Birth Order 3	0.259**	0.455**	0.589**	-0.025	-0.023	-0.061
Birth Order 4	0.130**	0.375**	0.416**	-0.157**	-0.176**	-0.170*
Birth Order 5 and higher	-0.129	-0.143	0.180	-0.186**	-0.176**	-0.170**
Number of vaccinations				0.068**	0.054*	0.095**
Previous child died or mother miscarried	-0.010	-0.106	0.008	0.061	0.150*	-0.044
Household demographic characteristics						
Mother is spouse of household head	-0.168**	-0.320**	-0.089	-0.016	-0.090	0.040
Proportion of daughters in the household	-0.134**	1.549**	-1.577**	0.119**	0.076	0.256**
Household size	0.016*	0.019	0.035*	-0.009	-0.015	-0.007
Number of children under 5	0.586**	0.721**	0.666**	-0.114**	-0.075*	-0.215**
Mother's weight	0.001	0.006	-0.003	0.022**	0.024**	0.020**
Mother's height	0.007	-0.005	0.015*	0.042**	0.034**	0.048**
Mother's education – primary	0.166**	0.304**	0.083	-0.072	-0.108*	-0.019
Mother's education – secondary	0.175**	0.267**	0.108	0.107**	0.109*	0.127*
Father's education –primary	0.020	0.001	0.020	0.067	0.152**	-0.030
Father's education – secondary	-0.001	-0.014	-0.011	0.056	0.122*	-0.029
Father's education – higher than secondary	0.289**	0.379*	0.165	0.254**	0.341**	0.139
Household economic characteristics						
Mother not working	0.040	0.225**	-0.076	-0.087	-0.164	0.013
Father's occupation – office worker	-0.030	-0.015	-0.019	-0.046	-0.033	-0.062
Father's occupation – manual labourer	-0.017	0.005	0.032	-0.057	-0.056	-0.056
Access to piped water for drinking	-0.068	-0.046	-0.211			
Wealth quintile – 2nd	0.188**	0.118	0.240**	0.130**	0.116	0.163*
Wealth quintile – 3rd	-0.041	-0.027	-0.104	0.280**	0.208**	0.354**
Wealth quintile – 4th	0.166**	0.081	0.170	0.274**	0.211**	0.341**
Wealth quintile – 5th	-0.034	-0.093	-0.004	0.627**	0.575**	0.693**
Mother's autonomy						
Where to go to access health care for self				-0.034	-0.113*	0.08
Needs permission to spend money				-0.098	-0.025	-0.154
Knows where to get help if child is ill				0.006	0.007	-0.003
Needs permission to access health care				0.084*	0.152**	0.014
Needs permission to spend money on health				-0.092**	-0.112**	-0.043
σ				1.190**	1.181**	1.180**
θ				-4.758**	-4.240**	-4.981**
Maximised log likelihood				-8693.24	-4188.41	-4348.21
Kendall's τ				-0.441**	-0.406**	-0.456**
N	5172	2533	2639	4711	2321	2390

Note: ** and * indicate significance at 1% and 5% levels respectively. Omitted, but available upon request, are estimates for child's age and age-squared, mother's age, father's age, father's age missing and regional variables.

Table 3: Maximum Likelihood Estimates: Weight-for-Height

	Stage 1- Child alive			Stage 2- weight-for-height estimates		
	Full sample	Female	Male	Full sample	Female	Male
Child characteristics						
Male	-0.123*			0.010		
Twin	-1.394**	-1.403**	-1.400**			
Birth Order 2	0.171**	0.394**	0.409**	-0.023	-0.115**	0.117**
Birth Order 3	0.237**	0.455**	0.511**	-0.023	-0.104*	0.112**
Birth Order 4	0.123	0.426**	0.311**	-0.044	-0.149**	0.068
Birth Order 5 and higher	-0.096	-0.093	0.161	0.035	-0.090	0.153**
Number of vaccinations				-0.023	-0.015	-0.026
Previous child died or mother miscarried	-0.003	-0.153	-0.016	0.092**	0.203**	0.017
Household demographic characteristics						
Mother is spouse of household head	-0.165**	-0.321**	-0.048	0.044	0.070	0.018
Proportion of daughters in the household	-0.116	1.526**	-1.592**	-0.026	-0.298**	-0.210**
Household size	0.025*	0.029	0.033*	0.015**	0.011	0.024**
Number of children under 5	0.580**	0.682**	0.673**	-0.062**	-0.058*	0.072**
Mother's weight	0.003	0.005	-0.002	0.025**	0.025**	0.026**
Mother's height	0.005	-0.001	0.014*	-0.005*	-0.011**	0.002
Mother's education – primary	0.165**	0.307**	0.144	0.046	0.038	0.066
Mother's education – secondary	0.190*	0.284*	0.178	0.067*	0.030	0.143**
Father's education – primary	0.030	0.012	-0.023	-0.037	0.002	-0.062
Father's education – secondary	0.061	0.040	0.031	-0.095**	-0.039	-0.161**
Father's education – higher than secondary	0.314**	0.432*	0.119	-0.042	0.030	-0.068
Household economic characteristics						
Mother not working	0.041	0.135*	-0.116*	-0.009	-0.078	0.049
Father's occupation – office worker	-0.069	-0.101	-0.073	0.053	0.107**	-0.012
Father's occupation – manual labourer	-0.031	-0.066	0.006	0.067**	0.112***	0.035
Household has piped water for drinking	-0.025	0.072	-0.090			
Wealth quintile – 2^{nd}	0.145*	0.058	0.250**	-0.004	-0.005	0.057
Wealth quintile – 3^{rd}	-0.015	0.007	-0.089	0.085**	0.056	0.116*
Wealth quintile – 4^{th}	0.117	0.054	0.134	0.011	0.022	0.043
Wealth quintile – 5^{th}	-0.081	-0.147	-0.076	0.032	0.030	0.026
Mother's autonomy						
Where to go to access health care for self				-0.029	-0.041	-0.002
Needs permission to spend money				-0.069	-0.143**	0.009
Knows where to get help if child is ill				0.142**	0.156**	0.138**
Needs permission to access health care				-0.086**	-0.109**	-0.053
Needs permission to spend money on health				0.127***	0.135**	0.103
σ				0.880**	0.869**	0.895**
θ				-5.661**	-6.526**	0.681**
Maximised log likelihood				-7232.2	-3438.2	-3633.09
Kendall's τ				-0.496**	-0.541**	0.477**
N	5172	2533	2639	4711	2321	2390

Note: ** and * indicate significance at 1% and 5% levels respectively. Omitted, but available upon request, are estimates for child's age and age-squared, mother's age, father's age, father's age missing and regional variables.

Figure 1: Non-parametric kernel density plot of height-for-age for children in the full sample of 'alive' group

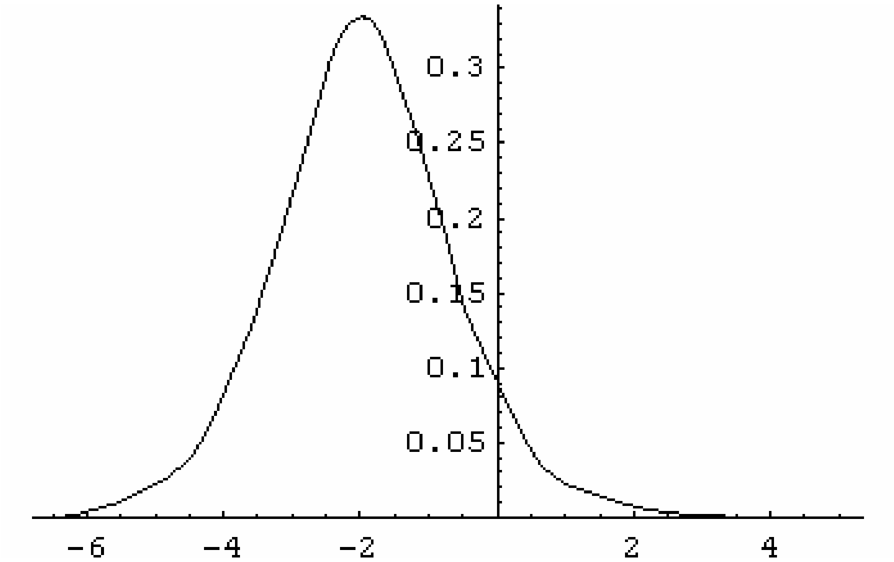

Figure 2: Non-parametric kernel density plot of weight-for-height for children in the full sample of 'alive' group