Does the number of sons born affect long-term mortality of parents? A cohort study in rural Bangladesh

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It has been suggested that bearing sons increases long-term mortality in women, because sons may be more physiologically demanding to produce than daughters. In this historical cohort study in rural Bangladesh, no association between the number of sons born and mortality was seen in women in the unadjusted analyses. However, a significant reduction in mortality with the number of surviving sons was seen. In addition, after adjusting for the number of surviving sons, there was evidence of increasing mortality with the number of sons born, in women. In men, mortality also depended strongly on the number of surviving sons, but not on the number born. These data provide support for negative long-term costs of bearing sons in mothers in rural Bangladesh, and suggest that there are context-specific factors that mask the true effects of sons in some populations.

Keywords: sons; boys; parents; mortality

1. INTRODUCTION

Evolutionary theories suggest that reproduction carries long-term costs because physical resources used during pregnancy cannot be used for later somatic repair (Kirkwood & Rose 1991). However, there is a little empirical evidence that mortality in women who have completed their reproduction increases with the number of births (for example, Hurt *et al.* in press). One proposed explanation for this apparent lack of an effect in humans is that, it may not be the number of children that is important but the cumulative number of sons born, because sons may be more physiologically demanding to produce than daughters (Helle *et al.* 2002; Mace *et al.* 2003).

There is evidence that women face additional physiological burdens when pregnant with sons. Sons are known to grow faster in the womb than daughters (Hindmarsh *et al.* 2002), and weigh more than daughters at birth (Marsal *et al.* 1996). The energy intake of women pregnant with sons is also higher than that of women carrying daughters, suggesting that they have higher energy requirements during pregnancy (Tamimi *et al.* 2003). In addition, the sex ratio at birth is biased towards daughters in some settings where mothers are malnourished, which has been interpreted as an evolutionary adaptation to protect women against the higher costs of bearing sons (Mace *et al.* 2003). Birth intervals in Kenya are longer after sons than daughters (Mace & Sear 1997)

and the likelihood of reproducing again is reduced after twin sons in Finland (Lummaa 2001).

Few studies have examined the association between the number of sons born and long-term mortality, and results have not been consistent. In Finland and rural Germany, the number of sons born significantly increased mortality in women age 50 and over, while there was a non-significant reduction in mortality in French migrants in Canada (Beise & Volland 2002; Helle *et al.* 2002). In an eighteenth century population in Flanders, no effect of sons was seen overall, but mortality significantly increased with the number of sons born, in women in the lowest socio-economic group (Van de Putte *et al.* 2004).

The interpretation of these results is complex. Biological and cultural factors may interact in different ways in different populations (Helle *et al.* 2002; Van de Putte *et al.* 2004), with a detrimental effect of the number of sons born masked where sons are also socially or economically beneficial to their parents. One way to determine whether biological or social factors predominate in any given population is to compare associations in women and men, because men are not subject to the physiological stresses of childbearing but may experience socio-economic benefits from having sons (Kravdal 1995). Others are to examine the effect of the number of sons born in different socio-economic strata (Van de Putte *et al.* 2004) or to separate the effects of the number of sons born from that of surviving sons, which no study has previously done.

The aim of this study was to assess whether long-term mortality increased with increasing number of sons born, in women and their husbands in rural Bangladesh. As surviving sons are thought to be advantageous to parents in this setting, we examined whether the effects of the

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number of sons was independent of, or modified by, the number of surviving sons. To further examine the role of social factors, we tested for an interaction with socio-economic status.

2. MATERIAL AND METHODS

(a) *Source population and data collection*

This study was conducted in Matlab, Bangladesh—a poor rural area southeast of the capital Dhaka. The population is predominantly Muslim, and the major sources of income are fishing and agricultural labour. During the study period, there was early and almost universal marriage and, within marriage, there was close to natural fertility (Stoeckel & Choudhury 1970; Razzaque *et al.* 1998).

This is a historical cohort study. We used data collected between 1982 and 1998 by the ICDDR, B: Centre for Health and Population Research (ICDDR, B) within a unique health and demographic surveillance system (HDSS). The HDSS covers the whole population of a delineated area in the Matlab district and provides longitudinal data on all births, deaths, marriages and migrations since 1966. Currently, 142 villages and approximately 200 000 individuals are under surveillance (Fauveau 1994). Literate female community health workers collect data on a monthly basis. They are resident in the area in which they work and cover a population of around 1000 individuals each.

(b) *Study cohort*

We examined all-cause mortality in women over the age of 45 and their husbands. Women entered the cohort if they were aged between 45 and 55 on census day (June 30, 1982), or if they became 45 or migrated into the area aged between 45 and 55 during the study period. We excluded younger women because their reproduction may not have been completed, and because early mortality may be a direct consequence of pregnancy. Women experiencing a pregnancy outcome in the year before entry or during the follow-up period did not contribute person-time until one year after this last pregnancy outcome, because deaths in this period may be a direct result of pregnancy (WHO 1992).

Husbands were included if they were alive and resident in the area on the day their wives entered the cohort, contributing person-time from the same day as their wives. There were fewer men than women, as 1985 women were widowed or divorced on entry and 3595 husbands were not resident in the area during the study period. Migration of men to work in Dhaka, India and the Middle East is common and they are not included in the data collection during these absences.

(c) *Statistical analyses*

Women were classified according to the number of sons born alive and the number surviving at the time they entered the cohort. It was assumed that the reproductive histories of the men were the same as their wife's. We also obtained data on age, marital status, parity, area of residence, migration and mortality from the HDSS. Data on religion, education and occupation were obtained from the 1982 census.

Person-years of follow-up were calculated from entry until the subject died, migrated or until December 31, 1998. The observations for the women known to be alive on the last day of the cohort were censored on that date. We calculated

unadjusted mortality rates with 95% confidence intervals (CI) by the number of sons born and surviving and used Poisson regression (Clayton & Hills 1993) to compare mortality rates at different exposure levels while adjusting for potential confounders.

Our strategy for the multivariate analysis was to first adjust for age, time period, marital and socio-economic status, keeping variables that had a significant impact (measured by a value $p < 0.10$ when using likelihood ratio tests) on the models. Then we added parity to assess whether any association between the number of sons and mortality was independent of the number of children. Then we adjusted for and stratified by the number of surviving sons, and examined for a statistical interaction between the number of sons born alive and the number surviving using likelihood ratio tests. The number of sons and number of surviving sons were examined as both continuous and categorical variables, and linear trends, were fitted only when there were no significant departures from linearity (measured using a likelihood ratio test). We also examined for an interaction between the number of sons born and socio-economic status. Lastly, we examined for an independent effect of the number of daughters and explored whether the effect of sons remained after adjustment for number of daughters.

There were missing reproductive histories in 1573 (8%) women and in 861 (6%) men. Excluding these individuals would mean making inappropriate assumptions about the unobserved data. We, therefore, used regression-based multiple imputation (Rubin 1987; Schafer 1997) to predict the missing values in our sample. All demographic and social variables were included in the predictive regression models and data on the outcome were also used, to avoid introducing bias into the imputed estimates (Schafer 1997). Imputations were performed separately for women and men, as the significant proportions of missing data on male education and occupation in the women made these variables inappropriate for use in the female imputations or subsequent analyses in the women. Five imputed datasets were obtained for each sex using SOLAS v. 2.0 (SOLAS 1999). Poisson regression was then performed in each dataset following the scheme outlined above, using STATA v.7.0 (Stata Corporation 2001) and the results combined using conventional multiple imputation rules (Schafer 1997).

3. RESULTS

The women had a mean age of 46.7 years (s.d.=2.8) on entry into the cohort (table 1). They had given birth to a mean of seven children (s.d.=2.6), with an average of 5.2 still alive on entry. There were few women (5.0%) who had not given birth to any sons, or had no surviving sons after age 45 (7.9%). The men were approximately 10 years older than their wives (mean age at entry 57.2, s.d.=5.7) and their fertility was somewhat higher (mean number of live births 7.3, s.d.=2.5). Although this difference was small, it was statistically significant ($z=10.93$, $p<0.001$). A slightly lower percentage had never had any sons (4.2%) or had no surviving sons on entry into the study (6.5%).

Unadjusted mortality rates in women reduced with the number of sons born until they had nine or more sons, after which they rose (figure 1). There was no statistically significant linear trend in these rates (unadjusted rate ratio (RR) per additional son born

Table 1. Distribution of baseline characteristics among women and their husbands, Matlab, 1982–1998. (s.d., standard deviation.)

	women			men		
	number of women	mean (number)	s.d. (%)	number of men	mean (number)	s.d. (%)
age at entry ^a	20 383	46.7	2.8	14 803	57.2	5.7
live births ^b	18 810	7.0	2.6	13 942	7.3	2.5
number of sons born						
0	18 810	1024	5.0	13 942	615	4.2
1–2		5094	25.0		3391	22.9
3–4		7424	36.4		5579	37.7
5–6		4129	20.3		3376	22.8
7–8		998	4.9		858	5.8
9+		141	0.7		123	0.8
number of surviving sons						
0	18 810	1615	7.9	13 942	966	6.5
1–2		7796	38.3		5351	36.2
3–4		7160	35.1		5682	38.4
5+		2239	11.0		1943	13.1

^a Data on age available for full cohort.

^b Data on reproductive history missing in 1573 (7.7%) of women and 861 (5.8%) of men, figures presented here are before imputation.

Figure 1. Unadjusted mortality rates per 1000 person-years (and 95% CI) by the number of sons, (a) born and (b) surviving in women.

0.98, 95% CI 0.96, 1.01; likelihood ratio test for trend $p=0.10$). However, mortality decreased with the number of surviving sons (unadjusted RR per additional surviving son 0.89, 95% CI 0.86, 0.90; test for trend $p<0.001$). In men, unadjusted mortality rates decreased significantly with the number of sons born and surviving (figure 2; unadjusted RR per son born 0.98, 95% CI 0.96, 0.99; test for trend $p=0.01$, and surviving 0.94, 95% CI 0.92, 0.96; test for trend $p<0.001$).

Adjusting for age, period, marital and socio-economic status did not change the patterns seen with the number of sons born, in women or men. There was no strong association between the total number of children born and mortality in women (adjusted RR per live birth 0.98, 95% CI 0.96, 1.01; test for trend $p=0.09$), but a significant decrease in mortality with the number of live births, in men (adjusted RR per live birth 0.97, 95% CI 0.96, 0.98; test for trend $p<0.01$). Adding parity to the models did not change the nature of the association between the number of sons born and mortality (see electronic supplementary material, tables A5 and A6). This variable was, therefore, not included in any further models.

Associations did, however, change when we adjusted for the number of surviving sons. In particular, the nature

of the patterns changed in women, and differed from those seen in the men. Mortality in women increased by 11% with every additional son born after adjusting for the number of surviving sons (table 2; 95% CI 1.06, 1.15; test for trend $p<0.01$). A significant protective effect of surviving sons also remained in this model (adjusted RR per surviving son 0.84, 95% CI 0.80, 0.88; test for trend $p<0.01$). There was no evidence of collinearity between the two variables (tolerance=0.319). There was also no evidence of a statistical interaction (LR test statistic 4.32; $p=0.31$), with the stratum-specific analyses showing a remarkably consistent effect of the number of boys born regardless of the number of surviving sons. The effect of the number of sons born did not modify with any of the other socio-economic variables.

Similar analyses looking at the effects of the number of daughters born showed no association (see electronic supplementary material, table A1). In addition, the effect of the number of sons born did not change when adjusted for the number of daughters or number of surviving daughters (see electronic supplementary material, table A3), and there was no interaction between the number of sons and the number of daughters.

In the men, there was no clear pattern in mortality with the number of sons born after adjusting or

Figure 2. Unadjusted mortality rates per 1000 person-years (and 95% CI) by the number of sons, (a) born and (b) surviving in men.

Table 2. Association between number of sons born and all-cause mortality in women age 45 and above, Matlab, 1982–1998: in full cohort and stratified by the number of surviving sons. (RR, rate ratio; CI, confidence intervals.)

characteristic	person-years ^a	mortality rate ^b	unadjusted RR	95% CI	adjusted RR ^c	95% CI
<i>full cohort</i>						
number of sons						
0	10 791	10.89	1.14	0.88, 1.39	0.58	0.39, 0.76
1–2	52 000	9.89	1.03	0.88, 1.18	0.76	0.64, 0.87
3–4	78 795	9.60	1.00	—	1.00	—
5–6	46 984	9.11	0.95	0.82, 1.08	1.11	0.95, 1.28
7–8	12 098	8.60	0.90	0.68, 1.12	1.16	0.84, 1.49
9+	1656	11.35	1.18	0.61, 1.76	1.73	0.80, 2.66
per son born	—	—	0.98	0.96, 1.01	1.11	1.06, 1.15
<i>0 surviving sons</i>						
number of sons						
0	10 791	10.90	0.65	0.32, 0.97	0.69	0.33, 1.05
1–2	5970	16.83	1.02	0.47, 1.57	0.95	0.41, 1.49
3+	1262	17.30	1.00	—	1.00	—
per son born	—	—	1.19	1.06, 1.32	1.15	1.01, 1.28
<i>1–2 surviving sons</i>						
number of sons						
1–2	46 029	8.94	0.70	0.59, 0.81	0.75	0.64, 0.87
3–4	30 573	12.47	1.00	—	1.00	—
5+	6466	11.75	0.90	0.66, 1.14	0.81	0.59, 1.02
per son born	—	—	1.10	1.03, 1.16	1.05	0.99, 1.12
<i>3–4 surviving sons</i>						
number of sons						
3–4	47 055	7.33	1.00	—	1.00	—
5–6	25 899	10.47	1.43	1.14, 1.73	1.32	1.04, 1.59
7+	3569	9.75	1.33	0.82, 1.85	1.12	0.69, 1.56
per son born	—	—	1.16	1.07, 1.25	1.10	1.01, 1.19
<i>5+ surviving sons</i>						
number of sons						
5–6	15 385	5.71	1.00	—	1.00	—
7–8	7951	8.75	1.54	0.99, 2.09	1.36	0.87, 1.85
9+	1371	9.77	1.71	0.69, 2.73	1.49	0.59, 2.38
per son born	—	—	1.20	1.04, 1.36	1.13	0.98, 1.29

^a Average over the five imputed datasets.

^b Per 1000 person-years of follow-up.

^c Full cohort adjusted for age, time period, religion, female education, area of residence, marital status and surviving sons; stratified results adjusted as for full cohort except for surviving sons.

stratifying for the number of surviving sons (table 3). Mortality decreased significantly with the number of surviving sons (adjusted RR per surviving son 0.94, 95% CI 0.91, 0.97, test for trend $p < 0.001$). There was no significant interaction between the number of sons born and the number surviving in men (LR test statistic 0.86; $p = 0.83$), and no effect of the number of daughters

born (see electronic supplementary material, tables A2, A4, A6).

4. DISCUSSION

Our results provide evidence that the number of sons born increases long-term mortality in women in rural

Table 3. Association between number of sons born and all-cause mortality in men age 45 and above, Matlab, 1982–1998: in full cohort and stratified by the number of surviving sons. (RR, rate ratio; CI, confidence intervals.)

characteristic	person-years ^a	mortality rate ^b	unadjusted RR	95% CI	adjusted RR ^c	95% CI
<i>full cohort</i>						
number of sons						
0	5723	38.51	1.09	0.94, 1.25	0.92	0.71, 1.14
1–2	27 956	37.09	1.05	0.97, 1.14	0.96	0.86, 1.06
3–4	48 575	35.22	1.00	—	1.00	—
5–6	32 241	34.07	0.97	0.89, 1.04	1.03	0.94, 1.13
7–8	8768	33.19	0.94	0.82, 1.06	1.02	0.87, 1.18
9+	1235	29.48	0.84	0.55, 1.12	1.11	0.68, 1.53
per son born	—	—	0.98	0.96, 0.99	1.02	0.98, 1.05
<i>0 surviving sons</i>						
number of sons						
0	5723	38.51	0.85	0.53, 1.16	1.00	0.62, 1.39
1–2	2841	42.30	0.93	0.55, 1.31	1.01	0.56, 1.45
3+	702	45.59	1.00	—	1.00	—
per son born	—	—	1.06	0.96, 1.16	1.00	0.90, 1.11
<i>1–2 surviving sons</i>						
number of sons						
1–2	25 115	36.50	0.94	0.84, 1.04	1.02	0.91, 1.12
3–4	17 285	38.76	1.00	—	1.00	—
5+	3596	40.22	1.03	0.85, 1.23	0.94	0.77, 1.12
per son born	—	—	1.02	0.99, 1.06	0.99	0.95, 1.02
<i>3–4 surviving sons</i>						
number of sons						
3–4	30 638	32.99	1.00	—	1.00	—
5–6	17 685	36.14	1.10	0.99, 1.20	1.03	0.92, 1.13
7+	2385	39.32	1.19	0.93, 1.45	1.00	0.77, 1.22
per son born	—	—	1.05	1.01, 1.09	1.00	0.96, 1.04
<i>5+ surviving sons</i>						
number of sons						
5–6	11 402	29.17	1.00	—	1.00	—
7–8	6054	30.45	1.04	0.85, 1.24	0.97	0.78, 1.15
9+	1072	27.26	0.93	0.57, 1.30	0.88	0.53, 1.22
per son born	—	—	1.01	0.93, 1.08	0.97	0.90, 1.05

^a Average over the five imputed datasets.^b Per 1000 person-years of follow-up.^c Full cohort adjusted for age, time period, religion, male education, male occupation, female education, area of residence, marital status and surviving sons; stratified results adjusted as for full cohort except for surviving sons.

Bangladesh. This effect was independent of the total number of children born, and only became apparent after accounting for surviving sons. In men, mortality depended strongly on the number of surviving sons, but not on the total number of sons born. No effect of the number of daughters born was seen in women or men.

Although the demographic data from the HDSS are likely to be accurate (Fauveau 1994), the validity of the results needs to be assessed. First, selection bias needs to be excluded as an explanation for the differences between women and men. There was no bias in the selection of women, but men who migrated were excluded and they may have differed in important ways from the men included in the study. The excluded men had slightly lower fertility than the general population. In addition, men who migrate are known to be better educated (Hadi 2001), but it is not known whether their mortality differs from those residing in Matlab. However, the patterns observed in women did not change when we restricted the analyses to those women whose husbands were present in the area, suggesting that the exclusion of migrant men

does not explain the differences in mortality between husbands and wives.

Second we used multiple imputation to accommodate the missing data. This technique makes less stringent assumptions about the unobserved data than other commonly used strategies such as exclusion, namely that the data are 'missing at random' (Rubin 1987; Schafer 1997). That is, the probability that the reproductive histories were missing was not related to the values of the reproductive histories once adjustment was made for other observed variables. There is no formal way to test this assumption, but we included factors that were associated with the missingness of the reproductive histories (such as education and migration) in the imputation models, making the assumption more plausible. On comparing the results from the imputed datasets with those from the dataset in which individuals with missing values were excluded, we obtained similar results. In addition, consistent results were obtained when we repeated the analyses in the sample of women whose husbands were present using data from the

male imputation models. Bias due to non-ignorable missingness is, therefore, unlikely to have had a major impact on the results.

There could be collinearity between the number of sons born and number surviving, as women must give birth to a certain number of sons to have a similar number surviving. However, there was no evidence of this. The correlation between the number of sons born and the number of sons surviving (which was 0.83) was lower than that usually thought to lead to collinearity (greater than 0.90). In addition, the unadjusted effect of each variable on the outcome was completely different (figures 1 and 2). Finally, the stratum-specific results presented are consistent with those adjusted for surviving sons, providing compelling evidence that the patterns seen are not an artefact of collinearity.

While we adjusted for the potential confounding effects of women's age, education, marital status, religion and area of residence, there may be additional woman-based confounders (in particular, those relating to frailty) for which we did not adjust. This may mean that we have underestimated the true effect of the number of boys, because women who have given birth to more boys may be inherently stronger and healthier than women with fewer successful male births, although the standard error of this effect may be larger than we have estimated.

Finally, there may be some misclassification of exposure. In this prospective cohort, any misclassification of the sex of children born is likely to be minimal (Becker & Mahmud 1984). The assumption that a man had the same reproductive history as his current wife is also reasonable, as polygamy and remarriage in men are rare and extramarital sexual activity is discouraged (Aziz & Mosley 1994; Razzaque *et al.* 1998). However, we underestimated the total number of sons conceived, because we only knew the sex of live births. If we assume that all sons conceived represent a physiological burden to women, it is possible that we underestimated their full effect due to this non-differential misclassification.

Our results are not consistent with those seen in women in Finland and Germany (Beise & Volland 2002; Helle *et al.* 2002), as these studies found a negative effect of sons without accounting for the number of surviving sons, but our unadjusted results are consistent with those seen in Canadian and Flemish women (Beise & Volland 2002; Van de Putte *et al.* 2004). We found no interaction with socio-economic variables such as education and religion, in contrast to the Flemish study (Van de Putte *et al.* 2004), but we did find that adjusting for another variable closely related to socio-economic status (the number of surviving sons) changed our results. In addition, there was no association between the number of sons and their fathers' mortality in Finland, whereas we saw a reduction in fathers' mortality until we accounted for the number of surviving sons. These differences lend support to the theory that there are context-specific factors (in our case, a strong protective effect of surviving sons) that mask a negative effect of sons on female mortality in some populations but not in others and shows how important it is to account for such factors. The lack of a negative effect in the men also supports the theory that there may be physiological costs of bearing sons that are specific to the mother.

It is plausible that there are higher evolutionary costs associated with producing sons, and that these may be enhanced in the chronically malnourished women of rural Bangladesh. However, there also remains a strong preference for sons in rural Bangladesh, with the number of surviving sons influencing both fertility behaviour and contraceptive use (Bairagi & Langsten 1986; Amin & Mariam 1987). Sons are valued because they contribute substantially to household income, establish and maintain family ties to village patronage groups, and continue the family name. They also provide care for elderly parents (Rahman 2000). It is plausible that sons surviving to adulthood reduce their parents' mortality through improvements in economic and social conditions (Mostafa & van-Ginneken 2000). These benefits seem particularly important in this setting and, overall, the positive effects of the number of surviving sons were stronger than the negative effect of sons born.

However, evolutionary theories also predict that all births, whatever their sex, should have an impact on long-term mortality. Although female pregnancies may have a short-term physiological impact, they do not seem to affect a woman's survival in the long-term in this context. This lends support to the theory that male and female pregnancies carry different costs, possibly because there is a physiological threshold (in terms of energy used, for example), which female pregnancies do not exceed but male pregnancies do.

Our results provide support for long-term negative costs to mothers of bearing sons in rural Bangladesh, and suggest that there are context-specific factors that mask this increased mortality in some populations but not in others. However, the interpretation of such results remains difficult and more research is required before the origin of these costs and the mechanisms by which they affect long-term mortality are fully understood.

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As this paper exceeds the maximum length normally permitted, the authors have agreed to contribute to production costs.