Son Preference and Fertility in China

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Abstract

This paper examines the effect of son preference on the hazards of having a second, and a third birth. With data from the Two-per-thousand National Sample Survey on Fertility and Contraception conducted in 1988 by the State Family Planning Commission of China, we examine the hazard of having a second birth among 62+ thousand married women who have had a first birth, and the hazard of having a third birth among 43+ thousand married women who have had two births. We analyze these two hazards (i.e., the hazard of moving from the 1st to the 2nd birth, and the hazard of moving from the 2nd to the 3rd birth) by estimating Cox proportional hazard models. The major co-variate in the first analysis is whether or not the first-born was a daughter. In the second analysis the main co-variate is whether both of the first two children were girls. In both models we control for seven co-variates known to have independent effects on the transition to a second (or third) birth, namely, whether the woman is a Han, whether she is a farmer, her age at the birth of the first (or second) child, whether she had her first (or second) birth prior to the initiation in 1979 of the one-child policy, and three dummy variables reflecting her level of education. Our results show the important influence of son preference on the hazard of having another birth.

SON PREFERENCE AND FERTILITY IN CHINA [1]

by

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INTRODUCTION

The true sex ratio at birth (SRB) does not vary much in the countries of the world. Countries with reasonable birth registration data report SRBs of around 104 to 107 male births for every 100 female births (Chahnazarian, 1991: 214). When the SRB is above or below the biological average range of 104-107, other factors, often sociological, such as sex preference, may well override the biological forces, biasing upwards, or downwards, the SRBs (Bumiller, 1990; Guttentag and Secord, 1983; among others).

In this paper we examine the impact of son preference on fertility in the People's Republic of China. We look specifically at the impact that a female birth has on the probability of a woman having a higher order birth. Our analysis should provide indirect information on one way son preference leads to a higher than average sex ratio at birth.

According to Zeng and his associates, the SRB in China “was very close to 106 in most years in the 1960s and 1970s” (1993: 283). Since the 1980s, particularly since the mid-1980s, the SRB has been on the increase, reaching 114.2 in 1992 (Gu and Roy, 1995: 20), and in 1994, the latest year for which we have data, the SRB reached 116.3 (Poston et al., 2000). However, the SRBs in China for parities above one are considerably higher than the SRB values in 1989 of 114 reported for all births, and of 105 for parity one births. To illustrate, the SRBs for parities 2 and 3 for the year of 1989 are 121 and 125, respectively; only the SRB values for parity one are in the normal range for most years.

Several answers have been advanced to the question about why the SRBs, particularly those above the first parity, are so abnormally high, none of which is inconsistent with an argument grounded in son preference. Some have reported that the increase in the reported sex ratios at birth in China is mainly due to female infanticide, or to prenatal sex identification followed by gender-specific induced abortion, or to the underreporting of female births (Hull, 1990), or to female out-adoption (Johansson & Nygren, 1991). Zeng and his associates (1993) hold that the sex-differential in the underreporting of births, along with sex selective abortion, explain almost all of the recent increase in the reported sex ratio at birth in China (see also Tuljapurkar et al., 1995). Surprisingly, few empirical inquiries entertain the influence of a less proximate influence such as son preference.

We endeavor in this paper to address this void. Specifically, among Chinese women with one birth, we will test statistically whether the presence of a daughter increases the probability of the women having a second birth. And among Chinese women with two births, we will ascertain whether having two daughters increases the probability of having a third birth. Before conducting the empirical inquiries, however, we review some of the literature on the sex ratio at birth, and the extent to which it is influenced by son preference.

PRIOR ANALYSES

Since the 1980s, the sex ratios at birth in China for parities above one have been higher than normal (Zeng et al., 1993; Tuljapurkar et al., 1995). These high sex ratios may well reflect the people's desire to attain a certain number of sons when they have fewer children. Wen (1993) has suggested that the influences of son preference interacting with a drastic fertility-control policy resulted in the high and rising overall SRB and the increasing sex ratios with parity. Li and Cooney (1993) assessed whether son preference actually affected compliance with China's one child policy; they found that the son preference was so pervasive that it discouraged compliance with the one child policy.

The direct causes for the high reported sex ratios at birth in China, as well as in South Korea and Taiwan, are the underreporting of female babies (female infants either living with their own families or informally adopted to others and not recorded as live births), a stopping rule in which fertility is terminated when a high number of sons is achieved, and, in a few cases, infanticide (Park, 1983; Hull, 1990; Johansson and Nygren, 1991; Zeng et al., 1993). In this paper we look beyond these immediate causes of high sex ratios toward less proximate causes. Son preference, along with socioeconomic development and gender inequality, among other factors, would appear to interact with the above mentioned factors to produce these high sex ratios at birth. As Gu has observed with respect to both South Korea and China, "sociocultural factors [such as son preference] will be ... more influential than economic" ones, at least for a time (Gu, 1994: 4).

The desire for many sons has been an integral part of Asian (and many other) societies for centuries. In many countries of the world today, couples usually want their first birth to be a son. The preference for sons in China, South Korea and Taiwan is “both pervasive and extreme [and] Chinese populations ... typically exhibit high levels of son preference as well” (Arnold and Liu, 1986: 223). In Bangladesh, South Korea, and Nepal, more than 95 percent of women with both a son and daughter who want another child typically want that child to be a boy (Arnold and Liu, 1986). Arnold and Liu suggest that son preference in China is the result of “deeply rooted Confucian traditions” and sons are desired for the purpose of “family propagation, old-age security, the provision of labor, and the performance of ancestral rites” (1986: 222). Many still hold to the old Chinese belief that “many sons [bring] much happiness” (Greenhalgh and Li, 1993: 10). Interestingly, those countries with among the strongest family planning programs are the ones with the highest levels of son preference (i.e. China, Taiwan, Republic of Korea, Hong Kong) (Birdsall, 1985).

In many countries, a woman’s value is stipulated in terms of how many sons she has born (Mitra, 1979). Bumiller reminds us of the popular Sanskrit blessing given to Hindu women at the time of their weddings: "May you be the mother of a hundred sons" (1990: 10). Also, Gu notes that Chinese women from both urban and rural areas believe they will be at risk of discrimination and insult if they do not produce a male child (Gu, 1994).

The existence of sex preference by itself, however, will not influence fertility and the sex ratio at birth "unless technologies that translate the preference to fertility modification are applied" (Park, 1994: 21). The evidence indicates that where such technologies as contraception, abortion, and the capability for diagnosing the sex of the fetus are present and available, son preference is likely to affect fertility behavior. These technologies are now available in China, justifying further our examination of the effects of son preference on fertility. We turn now to the empirical investigations.

DATA AND METHODS

Our objective is to ascertain whether having a female first birth increases the probability that a woman continues on to have a second birth; and if having daughters as the first two births increases a woman's probability to have a third child. If there is no son preference in the society, there should be no association between the sex of (an) already born child(ren) and the likelihood of having another. However, if parents had no preference for sons per se, or daughters per se, but did have a preference for one child of each sex, then it is likely they would be more inclined to have a third birth if their first two children were of the same sex. However, the probability of continuing should not be as high as in societies with a definite preference for sons.

There is a fair literature indicating that if parents already have one daughter, they will be more likely to continue childbearing than would be the case for parents with one son; and the converse has been found true for parents with one son (Pong, 1994; Park, 1994; DaVanzo and Starbird, 1991; Das, 1987; among others). For instance, Pong (1994) has written that having one or more sons "induces parents to adopt more effective or permanent methods of birth control ... As a result, mothers with more sons bear fewer additional children, or have a longer subsequent birth interval, compared with mothers with more daughters" (1994: 137-138).

In this paper we explore whether having one daughter increases the probability that a women will have another child. If the expected positive association is found, it will indicate one of the ways that son preference works to influence the SRB, namely, through an increase in the fertility of women with an already born daughter.

We use data from China's 2/1000 National Sample Survey of Fertility and Contraception (NSSFC) that was conducted in 1988 (State Family Planning Commission of China, 1990). The NSFFC collected pregnancy and contraceptive history for all married women in China between the ages of 15 and 57. We use a 15 percent national sample of the 2/1000 NSFFC. We undertake two analyses, the first of women with one child, and the second of women with two children. After imposing various restrictions on our samples (specified below), we have available for the first analysis a total of 60,213 married women who had at least one child by July, 1988, the date of the survey. We have available for the second analysis a total of 43,374 married women who had at least two children by July, 1988.

We undertake hazard analyses of women with one child ever-born, and of women with two children ever-born, to ascertain, first, whether or not the presence of a female first-born enhances the women's transition to a second birth; and, second, whether among women with two children having two daughters further increases the transition to a third birth.

The survival-time data for the two groups of Chinese women consist of two variables; one is a dummy variable indicating for each woman whether or not the event (the second birth for the first group of women, the third birth for the second group of women) occurred during the observation period; the second is a variable measuring the number of months that have elapsed since the last birth and the conception leading to the next birth or the censoring event. The dummy variable (CHILD1-2; or CHILD2-3) is coded 1 if the woman had a second (or a third) birth, and zero if otherwise. The second survival-time variable is an interval variable (MONTHS1-2; or MONTHS2-3) and reflects the number of months between the date of the last birth and the date of conception leading to the next birth, or between the date of the last birth and the date of the censoring event. The censoring events include such events as the woman reaches the end of the childbearing age (age 45); the woman has a pregnancy after the last birth that ends in a miscarriage or in a stillbirth or in an abortion; the woman becomes sterilized sometime after the last birth; and the date of the survey, July, 1988.

We first describe the survivor data for the two groups of women. The first group of women, those 60,213 married women with one birth, were at risk of having a second birth for a total of over two million (2,048,828) months, so that on average each woman had a duration of 34 months between the last birth and the conception leading to the next birth or the censoring event. Over 76 percent of these women had a second birth.

The second group of women are the 43,374 women with a second birth; over 66 percent of them had a third birth; they had a mean duration time of risk of having a third birth of over 52 months.

One way to describe the survival-time data for the two groups of women is to graph their Kaplan-Meier (K-M) survivor functions (Kaplan and Meier, 1958; Hamilton, 1998). For the first group of women, let $\mathbf{n_t}$ represent the number of women who have not given birth to a second child and are not censored at the beginning of time period $\mathbf{t}$; $\mathbf{d_t}$ represents the number of second children born to these women during time period $\mathbf{t}$. The formula (below) is the Kaplan-Meier estimator of surviving beyond time **t** (i.e., not having a second birth beyond time **t**), and is the product of survival probabilities in **t** and the preceding periods:

$$\mathbf{S(t)} = \prod_{j=t0}^{t} \{(n_j - d_j)/n_j\}$$

In Figure 1 we have graphed **S(t)** against the number of months between the first birth and the conception leading to the second birth. The K-M survivor curve for this first group of women shows the probabilities of surviving the hazard of having a second birth for each month of analysis time. The curve steps down rapidly from a probability of near 1.0 of surviving the hazard of a conception leading to a second birth just a few months after the birth of the first child, to a probability of around .15 by about the 100th month, leveling off by the 150th month to a probability of surviving having a second birth of about .11.

** FIGURE 1 ABOUT HERE **

Figure 2 is a plot of the K-M survivor curve for the second group of women, namely those who have two births. Their K-M curve also steps down rapidly from a survival probability of near 1.0 for the few months after the birth of their second child, to a probability of surviving the hazard of having a third birth of about .25 by around the 150th month, leveling out at this probability for the remainder of the risk period.

** FIGURE 2 ABOUT HERE **

We use Cox's partial-likelihood method to estimate continuous time proportional hazards models of the transitions from first to second birth, and from second to third births. The Cox proportional hazards model assumes the following form when all the independent variables (co-variates) are time-independent, that is, when their values do not change over time, as is the situation in both of our analyses:

$$\log h(t) = \log h_0(t) + b_1 x_1 + \ldots + b_k x_k$$

where $h_0(t)$ is an unspecified function of time $\mathbf{t}$, and x_1 to x_k are co-variates, and b_1 to b_k are parameters to be estimated. In our first analysis, the main co-variate of interest is a dummy variable indicating whether the first child is a daughter. In our second analysis it is a dummy variable indicating whether both of the first two children are daughters. One feature of the Cox model that makes it so attractive is that the function of time does not have to be specified.

The dependent variable, $\log h(t)$, is the hazard rate, which is an unobserved value gauging the instantaneous probability that a woman will have another birth during the interval since the previous birth (Allison, 1984; Yamiguchi, 1991).

Our general hypothesis is that among women with one child, there will be a positive association between the dummy variable measuring whether or not the first child is a daughter and the instantaneous hazard of having a second birth, controlling for several other independent variables. In the second analysis we expect a positive association between the dummy variable indicating that both of the children are girls and the hazard of having a third birth. These associations should be positive owing to a preference among parents for sons over daughters. This association is based on the expectation that women will be showing son preference if they are more prone to transition to another birth if an earlier-born child is a girl and not a boy (Pong, 1994: 141). If there is no or only minimal sex preference in the society, the sex of the prior-born child(ren) should not be associated with the hazard of having another birth.

We made some restrictions to our sample data. A woman whose baby died more than nine months before her next pregnancy was not counted as a live birth. We do not expect that the sex of this now dead child will have an impact on the woman's probability of having another child. However, a child who died less than 9 months before the termination of the woman’s next pregnancy is considered as a live birth.

Also, the first birth (or the second birth in our second analysis) needs to have occurred at least 9 months before the date of survey (July, 1988). A woman who had her birth less than 9 months before the date of the survey had virtually no chance of having another birth by the date of the survey. Such a woman would not be included in our study.

Furthermore, the number of months between the first birth and the second birth should not equal zero. Whenever this occurred, we assumed that the woman had twin births, so we excluded her from our study. Our assumption is that the sex of the first twin birth has no impact on the probability of having the second twin birth. Women with twins were also excluded from our second analysis.

Our first analysis examines the hazard of moving from a first to a second birth. The main independent variable in this analysis is a dummy variable (GIRL1), scored 1 if the first child is a girl, and 0 if a boy. In the second analysis we examine the hazard of moving from a second to a third birth. In this analysis we include three dummy variables reflecting the sex composition of the woman's first two children. One dummy variable (GIRLS) is scored 1 if both of the first two children are girls, 0 if otherwise; a second dummy variable (DAUGHSON) is scored 1 if the first child is a girl and the second a boy, 0 if otherwise; a third dummy variable (SONDAUGH) is scored 1 if the first child is a boy and the second a girl, 0 if otherwise; in this equation, mothers with two boy children are the reference group. We expect that if the first two births are both girls, then the mother will have a high positive probability to go on to a third birth.

In both the first and second analyses we use seven additional co-variates as control variables, as follows:

(1) whether the woman is a member of the HAN majority nationality, scored 1 if yes, 0 if no;

(2) whether the woman is a FARMER (i.e., engaged in agricultural labor), scored 1 if yes, 0 if no;

(3) the woman's age when she had her first child (AGE-BIRTH1), or her second child (AGE-BIRTH2), measured in years;

(4) whether the woman had her first child (or her second child) by July, 1979, the date of the initiation of the one-child policy (POLICY-1 or POLICY-2), scored 1 if yes, 0 if no.

The last three co-variates are dummy variables representing the woman's amount of education, namely, (5) whether she completed elementary school (6 years of school) (ELEMENTARY), scored 1 if yes, 0 if no; (6) whether she completed junior high school (9 years of school) (JUNIOR), scored 1 if yes, 0 if no; and (7) whether she completed high school or more (12+ years of school) (HIGH), scored 1 if yes, 0 if no. Women who have completed less then elementary school comprise the reference group.

Among the 60+ thousand Chinese women with one live birth, 76 percent of them report having a second birth (CHILD1-2). The first births of 51 percent of these women were daughters (GIRL1). Also, 90 percent of these women are Han, and 67 percent are agricultural workers (FARMER). On average they were 23 years of age when they had their first births (AGE-BIRTH1), and 61 percent of them had their first babies before the initiation of the one-child policy in 1979 (POLICY-1). Regarding their educational attainment, 27 percent have an elementary education (6 years of school), 20 percent have a junior high education (9 years of school), and 11 percent have a high school education or more (12+ years of school). This means that the rest of the women, or about 42 percent of them, have less than an elementary education, and are thus considered as semi-literate or illiterate.

Among the more than 43 thousand married Chinese women with two children, 66 percent of them report having a third birth (CHILD2-3). Regarding the sex composition of the women's first two births, 24 percent had two girls (GIRLS), 26 percent had a girl and then a boy (DAUGHSON), 25 percent had a boy and then a girl (SONDAUGH), and 25 percent had two boys (BOYS). These women were, on average, 25 years of age when they had their second babies (AGE-BIRTH2), and 70 percent of them had their second babies before the initiation of the one-child policy in 1979 (POLICY-2). The values of the rest of the control co-variates are similar to those for the first group of women.

We use these seven co-variates as controls because of their independent effects on fertility. For instance, Han women typically have fewer children than non-Han women

(Poston and Shu, 1987), and agricultural women usually have more children than non-agricultural women (Poston and Gu, 1987; Poston and Jia, 1990). The later a woman's age at the time of her first birth, the fewer the number of children she should have ever-born to her. Also, given the requirements of China's one-child policy (Poston, 1986), if a woman had her first child before the start of the one-child policy, she would be more likely to go on to have a second child. Finally, the higher the woman's level of education, the fewer her number of ever-born children (Freedman et al., 1988; Poston and Gu, 1987).

We turn now to the results of our analyses.

RESULTS

Table 1 reports the Cox proportional hazard estimates of the effect of the presence of a daughter, along with the control co-variates, on the hazard of having a second birth for those 60+ thousand married women with one live birth. The most important result in Table 1 is the positive hazard coefficient for the GIRL1 variable. Among the married Chinese women in the sample this variable has a coefficient of .17 (see the top row of the first column of data in Table 1). This means that women whose first child was a daughter have a significant probability of going on to have a second birth. Recall that this positive and significant hazard coefficient of 0.17 is net of the effects on the probability of having a second birth of the seven control co-variates of whether the woman is a Han, whether she is a farmer, her age at the birth of the first child, whether she had her first birth prior to the initiation of the one-child policy, and her level of education. This means that among married women in China who have one child, they are significantly more likely to experience the hazard of having a second birth if their first birth was a girl than if it was a boy. Such a finding would not be expected in a society with no or very minimal son preference.

** TABLE 1 ABOUT HERE **

If we exponentiate the values of the hazard coefficients, that is, take their antilogs, we get hazard ratios (these are shown in the third column of data in Table 1). These exponentiated coefficients may be interpreted as follows: for each unit increase in the co-variate, the hazard is multiplied by its exponentiated coefficient. Thus, if we compute **$100(e^{b}-1)$** we get the percentage change in the hazard with each one unit change in the explanatory variable.

GIRL1 has a hazard ratio among the Chinese women of 1.18. This means that among married Chinese women, having a daughter as the first child instead of a son increases the hazard of having a second birth by 18%, or [100 (1.18 - 1 = .18)] = 18. The hazard coefficient and hazard ratio for GIRL1 indicate a sizable and significant positive effect of having a daughter on the hazard, or probability, of experiencing the transition to a second birth.

The results in the Cox model of the control co-variates confirm in many instances what we already know about their effects on the probability of having a second (or higher-order) birth. We would expect that since members of most of China's ethnic minorities have higher fertility than the Han majority (Poston and Shu, 1987), Han women would be less likely than minority women to have a second child. The hazard ratio for HAN of .91 means that the hazard for Han women of having a second birth is 9 percent less than that for non-Han women. Conversely, as expected, the hazard of having a second birth for farming women is 52 percent higher than that for non-farming women.

The older the woman when she has her first birth, the less likely she will experience the hazard of having a second birth. For every additional one year of age at the birth of her first child, the hazard of having a second birth is reduced by 5 percent. (A reviewer of an earlier version of this paper asked that we test the quadratic effect of age, by including age squared in the equation, because the effect of age on fecundity is non-linear. We estimated another hazard equation, including the co-variate of age squared, but its hazard coefficient was not significantly different from zero.)

Regarding the effects of the one-child policy, if a woman had her first birth prior to 1979, the hazard of her having a second birth is 75 percent higher than if she had her first birth after the start of the one-child policy.

Finally, we expected that increases in education should be negatively associated with the probability of experiencing the transition to a second child; this hypothesis is pretty much supported. The hazard of having a second birth for women who have completed junior high school is 22 percent less than that for illiterate and semi-literate women. For women who have completed high school or more, the hazard is 38 percent less.

A reviewer of an earlier version of this paper asked about the relative impacts of the co-variates on the hazard of having a second birth. We are not really as concerned with whether the GIRL1 variable is the most important predictor, but whether it is a statistically significant predictor, net of the effects of the other co-variates. Nevertheless, we will address the reviewer's important question. One way of assessing the relative impacts of the co-variates on the hazard is to raise the hazard ratio of each co-variate to the power of one standard deviation (Rabe-Hesketh and Everitt, 2000: 155). We have produced such semi-standardized hazard ratios and present them in the last column of data in Table 1. Although there is a problem in the interpretation of the meaning of semi-standardized hazard ratios when the co-variate is a dummy variable (cf. Long, 1997), their values nevertheless indicate the relative effects of the co-variates on the hazard of having a second birth.

The semi-standardized hazard ratios indicate that the most influential co-variates are those pertaining to "farmer" status and whether the first baby was born after 1979, the date of the initiation of the one-child policy. Of the eight co-variates, the GIRL1 variable is the fifth most influential.

We turn now to the results of the Cox proportional hazard of having a third birth for those 43+ thousand married women with two live births. As already noted, we will gauge the effect of son preference on the transition to the third birth by introducing three co-variates pertaining to the gender mix of the first two children, as follows: one dummy variable (GIRLS) is scored 1 if both of the first two children were girls, 0 if otherwise; a second dummy variable (DAUGHSON) is scored 1 if the first child was a girl and the second was a boy, 0 if otherwise; a third dummy variable (SONDAUGH) is scored 1 if the first child was a boy and the second was a girl, 0 if otherwise; mothers with two boy children are the reference group. We expect that if the first two births are both girls, then the mother will have a high positive probability to go on to a third birth. The effects of a balanced gender mix in the first two children on the hazard of having a third birth should be appreciably lower because one male baby would have been already born. These three dummy variables are included in the hazard equation along with the other seven co-variates that we used as control variables in the first equation.

Table 2 shows the results of this Cox analysis of the transition to a third birth. The hazard coefficient of GIRLS is 0.45. Women whose first two births are girls have a hazard of moving to a third birth that is 57 percent higher than that of women whose first two births were both boys. The hazard coefficients for women whose first two children were one of each sex were considerably less.

** TABLE 2 ABOUT HERE **

This finding of a hazard coefficient of 0.45 for the GIRLS co-variate in Table 2 may be the most striking instance of the effect of sex preference on fertility of the analyses reported in this paper. The Cox results indicate that if the woman’s first two children were first a girl and then a boy (DAUGHSON), the hazard of moving to a third birth was not statistically different from zero. If the woman’s first two children were first a boy and then a girl (SONDAUGH), the hazard of moving to a third birth was about 5 percent greater than if the first two births were both boys. However, as mentioned, if the first two births were both girls, her probability of having a third birth was 57 percent greater than if her first two births were both boys.

This positive and very high hazard coefficient for the GIRLS variable (Table 2) indicates well the normative requirements in Chinese society for women to bear one or more sons. They are much more likely to stop, or to delay, their childbearing once they have the son. But if the first two children are both girls, they are very likely to go on to have a third child.

What about the effects of the other co-variates, namely, those not dealing with the gender mix of the first two children? The effects of these co-variates are just as expected. Han women have a hazard of having a third child that is 17 percent less than that of non-Han women. Farming women have a hazard 18 percent higher than non-farming women. The older the woman's age at her second birth, the less likely she is to have a third birth. If a woman had her second birth prior to 1979, the hazard of her having a third birth is 132 percent higher than if she had her second birth after the start of the one-child policy. All three of the education dummy variables show the expected negative hazards of having a third birth, compared to illiterate and semi-literate women.

We also produced semi-standardized hazard ratios, and they are shown in the last column of Table 2. These indicate the relative effects of the co-variates on the hazard of having a third birth. The co-variate with the greatest relative effect on the hazard of having a third birth is whether the second baby was born before 1979 (POLICY-2), the date of the initiation of the one-child policy. The co-variate with the second most influential relative effect on the hazard of having a third birth is whether the woman's first two births were girls (GIRLS). If a woman's first two births are both girls, there appears to be an extraordinary amount of pressure for her to have a third birth; she must move to that third birth for another chance of having the prized son.

What are some of the implications of these inquiries? We address this issue in the last section of the paper.

IMPLICATIONS

This paper has demonstrated strong negative associations between son preference and fertility. Among women with one birth, those with a daughter are considerably more likely to have another birth, compared to those with a son. Among women with two births, those whose first two births are both girls are much more likely to have a third birth than if one or both of their first two children were boys. In this final section we link our results on the effects of sex preference on fertility with our earlier discussion about abnormally high sex ratios at birth (SRBs). We hold that abnormally high SRBs would likely not be occurring were it not for normative pressures in the society for sons.

Son preference patterns and effects such as those shown here, however, are not unique to China. In other societies, viz., Taiwan and South Korea, with rapid fertility decline and strong son preference, the normative requirement for sons is also being manifested in abnormally high SRBs at parities beyond the first.

It could be the case that the abnormally high SRBs now characterizing China, Taiwan and South Korea may be a transitory phenomenon that will diminish as these societies move to full modernization. The problem of the high sex ratio at birth, perhaps, can best be addressed when gender inequities are narrowed and strong son preference reduced; these kinds of changes are most likely to occur with increases in socioeconomic development and modernization (cf., Poston and Gu, 1987).

It is also true that the abnormally high SRBs in China may well be associated with the one-child policy instituted in the country in 1979 (Li and Cooney, 1993). But this does not necessarily mean that China would not be plagued by high SRBs if the one-child policy were relaxed or even terminated. Witness the high sex ratios in Taiwan and South

Korea where such drastic fertility policies do not exist (Poston et al., 2000; Poston et al., 2001). Were China's policy relaxed, the main difference would likely be that the SRB deviation from normality would not occur until after the second parity, instead of, as is now the case in China, after the first parity.

In an important manner, the high SRBs in China, and also in Taiwan and South Korea, have much to do with the rapidity of the fertility decline and the concomitant downsizing of norms concerning family size. However, the evidence from other Asian countries experiencing rapid fertility declines, namely Thailand, Sri Lanka and Indonesia, indicates that abnormally high SRBs are not occurring there. If son preference is low, as it is in these three South and Southeast Asia countries, rapid fertility declines will not lead to abnormally high SRBs.

Discussions in this paper, and elsewhere (Ng and Gu, 1995; Gu and Li, 1994), point to the importance of viewing the fertility transition in China and in other countries, not as a single dimension process focusing only on changes in fertility. Rather the fertility transition needs to be regarded as an integral component of the overall transition of the society from a traditional form to a modern one. One needs to consider not only the level of fertility, but also the timing of childbearing and the sex composition of the children. When fertility declines rapidly with regard to the number of children a couple may have, the tradition of strong preference for sons over daughters will become more salient; accordingly, socio-cultural factors will likely be more influential than economic ones in the fertility decision process. This may be why the patterns and trends of abnormally high SRBs observed in China are also seen in some other Asian countries with similar cultural contexts and rapid fertility declines.

The phenomenon of abnormally high SRBs occurring along with rapid fertility declines has only begun to receive the attention it deserves. Systematic field investigations combining both quantitative and qualitative methods are needed to help improve our understanding of the various social, economic, and cultural factors that determine people's son preference in childbearing, particularly in the low fertility settings now characterizing many Asian countries. These analyses will not only be important for the countries themselves, but they may also be of aid and relevance for other countries which may encounter similar problems at a later time.

ENDNOTE

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Table 1.
Cox Proportional Hazard Model Estimates of
the Effect of the Presence of a Daughter, and other Co-variates,
on the Hazard of Having a Second Birth:
60,213 Married Women, Aged 15-45, With One Live Birth

Variable	Hazard Coefficient	Z-score	Hazard Ratio	Semi-Standardized Hazard Ratio
GIRL1	0.17*	17.74	1.18	1.09
HAN	-0.09*	-6.18	0.91	0.97
FARMER	0.42*	33.98	1.52	1.22
AGE-BIRTH1	-0.05*	-32.98	0.95	0.85
POLICY-1	0.56*	47.04	1.75	1.31
ELEMENTARY	0.05*	4.81	1.05	1.02
JUNIOR	-0.25*	-15.90	0.78	0.91
HIGH	-0.48*	-20.56	0.62	0.86

Final Log Likelihood	=	-464706.45
Likelihood Ratio $\chi 2$	=	12592.71*

*significant at $p<0.05$, one-tailed test

Table 2.
Cox Proportional Hazard Model Estimates of
the Effect of the Presence of Two Daughters, and other Co-variates,
on the Hazard of Having a Third Birth:
43,374 Married Women, Aged 15-45, With Two Live Births

Variable	Hazard Coefficient	Z-score	Hazard Ratio	Semi-Standardized Hazard Ratio
GIRLS	0.45*	27.22	1.57	1.21
DAUGHSON	-0.00	-0.14	1.00	1.00
SONDAUGH	0.05*	3.10	1.05	1.02
HAN	-0.31*	-17.17	0.73	0.91
FARMER	0.17*	11.09	1.18	1.08
AGE-BIRTH2	-0.08*	-41.99	0.92	0.75
POLICY-2	0.84*	48.38	2.32	1.47
ELEMENTARY	-0.19*	-14.00	0.82	0.91
JUNIOR	-0.45*	-20.24	0.64	0.86
HIGH	-0.68*	-18.44	0.51	0.86

Final Log Likelihood = -284631.81
Likelihood Ratio $\chi 2$ = 8947.13*

*significant at p<0.05, one-tailed test

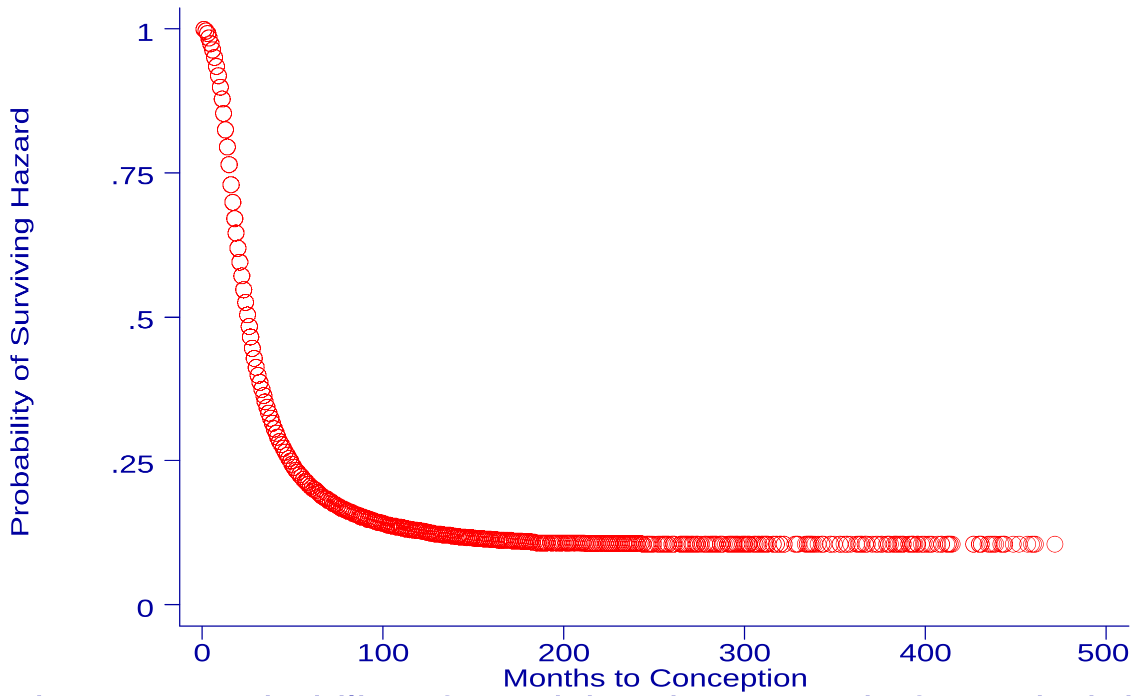

Figure 1. Probability of Surviving the Hazard of a 2nd Birth

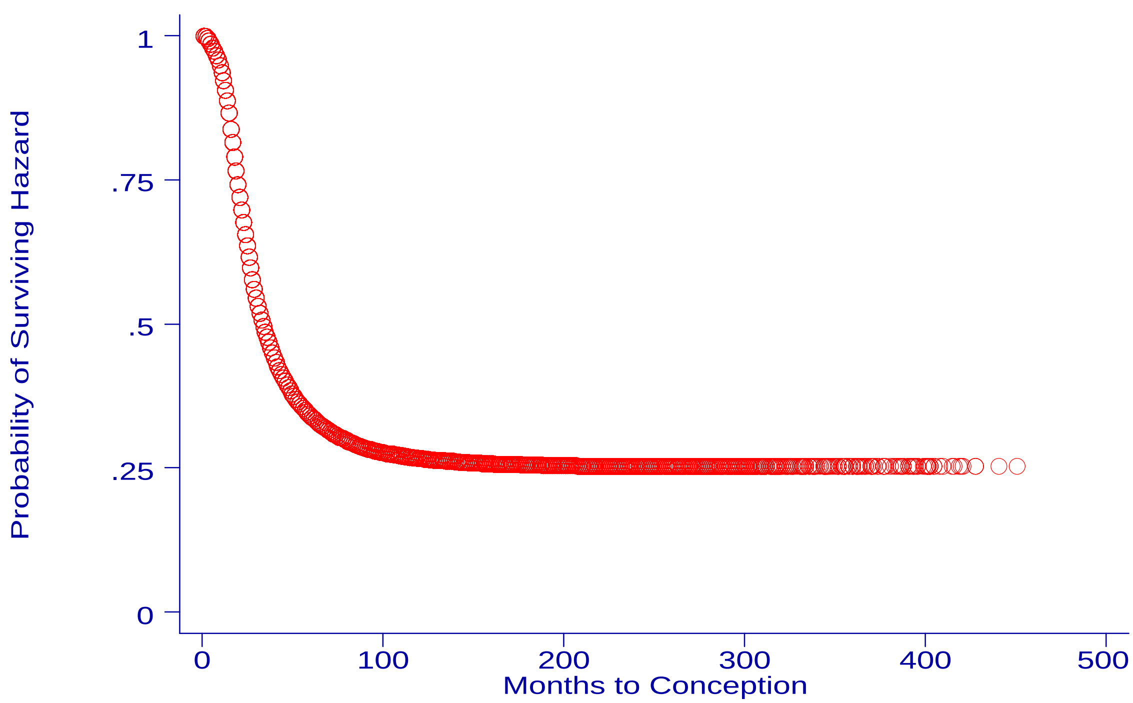

Figure 2. Probability of Surviving the Hazard of a 3rd Birth