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Equilibrating the Marriage Market in a Rapidly  
Growing Population: Evidence from Rural Bangladesh

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## Abstract

In this paper we show how relatively small changes in the age at marriage can equilibrate the marriage market despite relatively large differences in the supply of men and women of marriageable age. In particular, a simple demographic model is developed that provides a prediction about the relationship between relative cohort size and age at marriage that differs markedly from that of the static model that has been emphasized in the previous literature. We illustrate this point empirically using micro-level data from a rural area of Bangladesh which experienced a large increase in age at marriage for women (3.2 years) over the period from 1975 to 1990 at the same time that the relative supply of women in the marriage market was decreasing. We then estimate a behavioral model characterizing the relative values of men and women with different characteristics in the marriage market in order to better understand why female age at marriage served as the primary equilibrating mechanism in this population. The estimates suggest that the rise in the age at marriage was in part facilitated by a decline in the extent to which youth was valued by potential grooms and/or their families.

## I. Introduction

Changes in the age at marriage, particularly for women, have long been thought to be an important component of the development process. By delaying marriage for women, it is argued, women may stay in school longer, find more suitable mates, have greater say in the allocation of household resources, and begin childbearing at a later age which, in turn, may improve outcomes for children, result in fewer overall births per women, and slow population growth<sup>1</sup> In response to these concerns policies have been implemented that are designed to delay marriage directly through regulation and indirectly through financial incentives.<sup>2</sup> There is also substantial concern about potential social implications of an imbalanced number of men and women who wish to marry at a particular point in time (a “marriage squeeze”) and how this may influence and be influenced by the age at marriage. An excess supply of women that arises from a large gap in ages at marriages may increase marital payments from women’s to men’s families (Caldwell, Reddy and Caldwell 1983, Lindenbaum 1981, Rao 1993) and thus may be a factor in, for example, excess rates of female mortality. By contrast, there is concern that differential mortality and sex-selective abortion may lead to a substantial fraction of unmarried men and that changes in age at marriage can do little to address this imbalance (Tuljapurkar et al 1995).

Despite the apparent importance of age at marriage in relation to other individual and social measures of well-being, relatively little is understood about the underlying mechanisms of age at marriage change. While there is a useful theoretical literature on matching in general and

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<sup>1</sup>Singh and Samara (1996) summarize these arguments.

<sup>2</sup>In China, for example, the legal minimum age of marriage for women was increased to 20 in 1980 as part of an effort to lower fertility. Harayana State in India experimented with a program to provide bonds to low income families at the birth of a daughter that could be redeemed at age 18 only if the girl had not yet married (Singh and Samara 1996)

marriage markets in particular,<sup>3</sup> few theories have been examined empirically. Most studies that have examined age at marriage are partial-equilibrium in nature and thus miss the important general-equilibrium effects in the marriage market. Boulier and Rosenzweig (1984) show, for example, that less attractive women tend to marry later, that increases in schooling tend to decrease search time, and that longer search and thus later age at marriage, *ceteris paribus*, improves the quality of spouses, it is not obvious what these results imply about the effects of changes in average schooling on average marriage age or how delays in marriage as a whole affect equilibrium in the marriage market. In contrast to other demographic outcomes, which primarily involve behaviors of a single household, marriage involves interaction among households and thus inter-household and market-level relationships are likely to be important. Unfortunately, few data sets contain sufficient information on potential marriage partners to examine anything but the simplest forms of interhousehold interaction.

One of the few market-level theories of age at marriage-change that has been subject to substantial empirical scrutiny involves the relationship between population age distribution and the marriage squeeze. The basic idea, which obtains in a simple static model, is that the extent of the marriage squeeze associated with variation in cohort size will be smaller the smaller is the gap in age at marriage between men and women. In the extreme, given that sex ratios at birth are close to one, no marriage-squeeze arises at all when the age-gap is zero, regardless of variation in cohort size. Thus one would expect smaller gaps in age at marriage when the cohorts of women at peak ages of female marriage are substantially larger than the cohorts of men at peak ages of male marriage.

Despite the apparent simplicity of the posited relationship, the evidence is only weakly

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<sup>3</sup>Bergstrom (1999) provides an excellent summary of the relevant literature.

supportive, with most studies finding significant but weak relationships of the expected sign and others finding contradictory results (Smith 1983). Results from time series studies appear to be in general stronger than those from cross-sectional studies. For example, Bergstrom and Lam (1993), in an analysis of Swedish data, find that the difference in the ages at marriage between men and women responds to the relative number of men and women eligible for marriage as predicted by this static model.<sup>4</sup> Edlund (1999), however, using cross-sectional aggregate data from Asia uncovers only a weak relationship between sex-ratios and the spousal age gap.

In this paper we argue that the standard prediction of the cross-sectional model can be misleading because it does not account for key dynamic aspects of the process of marriage-market equilibration through changes in the age at marriage. In particular, a simple demographic model is used to show that the relevant relationship is not between relative cohort size and the *amount* of the age gap, but between relative cohort size and the *rate of change* in the age gap. We illustrate this point empirically using micro-level data from a rural area of Bangladesh which experienced a marked increase in age at marriage for women (3.2 years) over the period from 1975 to 1990 at the same time that the relative supply of women in the marriage market was not decreasing. We show that this result, which is at odds with the predictions of the static model, conforms well with the predictions of the alternative model. We then estimate a behavioral model characterizing the relative values of men and women with different characteristics in the marriage market in order to better understand why female age at marriage served as the primary equilibrating mechanism in this population. The estimates suggest that the rise in the age at marriage was in part facilitated by a decline in the extent to which youth was valued by potential

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<sup>4</sup>Note that the theoretical model provided by Bergstrom and Lam (1993) is dynamic and could, in principle, have yielded a test similar to the one developed below; the primary implication they test, however, is the standard one that arises from a static model.

grooms and/or their families.

## II. Demographic Model

The starting point for our analysis is a model of equilibrium that incorporates the idea of demographic translation, which is a tool used by demographers to relate cohort and period summary measures of age-specific demographic data (Ryder 1964, Foster 1990). Its most prominent application, to date, is in the context of the US baby boom where it was recognized that the substantial increases in average rates of childbearing during the 1950s (period fertility) were not matched by corresponding increases in completed family size (cohort fertility) by women in the peak childbearing averages during these years. Demographic translation was used to show that this discrepancy could be largely accounted for by the fact that there was a substantial drop across cohorts in the average age at childbearing so that previous cohorts who had delayed their fertility were giving birth at the same time as subsequent cohorts with earlier childbearing (Ryder 1964).

In the current context demographic translation is used to convert cohort and sex-specific measures of cohort size in combination with the timing and rate of marriage into period-specific counts of the number of men and women wishing to marry at each point in time. Thus, let  $\phi_m(x, T)$  denote the proportion of males born in year  $T$  who marry at age  $x$  (inclusive of remarriage),<sup>5</sup> and let  $N_m(T)$  denote the number of males in cohort  $T$ . Further, let  $c_m(T) = (N_m(T) - N)/N$  denote variation in cohort size relative to the average cohort, which is assumed to be of size  $N$ . Then by integrating across all cohorts alive at time  $t$  scaled by the appropriate age-specific

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<sup>5</sup>Note that  $\phi_m(x)$  is the marital analog to the net maternity schedule: it is the product of the probability of marriage at age  $x$  among survivors and the probability of surviving to age  $x$ . It also permits more than one marriage for each individual. The reason for examining this schedule rather than the more traditional first-marriage hazard is the need to keep track of all marriages in the economy including, for example, second marriages of men to first marriages of women.

marriage rate we may write the number of males marrying at period  $t$  as:

$$M_m(t) = N \int_0^{\omega} (1 + c_m(t-x)) \phi_m(x, t-x) dx \quad (1)$$

where  $\omega$  is the maximum age in the population. A similar expression represents the number of women marrying in period  $t$ ,  $M_f(t)$ , with each subscript  $m$  replaced with an  $f$ . Marriage-market equilibrium requires  $M_m(t) = M_f(t)$  at each point in time.

We then parameterize the cohort-specific marriage rate schedule  $\phi_m(x, T)$ . We rely on a two-parameter model<sup>6</sup> which related the marriage rate schedule in a particular cohort  $T$  to the age-pattern of marriage in an average or standard population,  $\phi_{ms}(x)$ <sup>7</sup> The two parameters of the model,  $g_m(T)$  and  $a_m(T)$ , as illustrated below, characterize the overall frequency of marriage for cohort  $T$  and the age-distribution or timing of marriage, respectively, and are written as functions of  $T$  to emphasize the fact that the parameters will, in general, differ by cohort. In particular, the cohort-specific rate of marriage for someone of age  $x$  in cohort  $T$  is modeled:

$$\phi_{ms}(x, T) = (1 + g_m(T)) \phi_{ms}(x - a_m(T)) \quad (2)$$

We also assume that the standard schedule  $\phi_{ms}(x)$  is continuous and differentiable in  $x$ , that

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<sup>6</sup>In the field of demography this is known as a model schedule. Demographic model schedules take advantage of the substantial regularity in age patterns of demographic rates and are primarily used to construct age-specific vital rates using incomplete data. For example, in a one-parameter model information on child mortality might be used to predict the whole age-pattern of mortality in a given population. Formally, a  $k$ -parameter demographic model schedule is a mapping  $\phi: [0, \omega] \times \mathfrak{R}^k \rightarrow \mathfrak{R}$  that yields for each parameter value a predicted vital rate at each age. The particular schedule created here is a simplification of that developed in Foster (1990).

<sup>7</sup>The choice of a standardized schedule is, of course, somewhat arbitrary but one reasonable approach, which is followed in this paper, is to average data on age-specific rates across time.

there is a minimum  $\alpha$  and maximum  $\beta$  age of marriage such that  $\phi_{ms}(x)=0$  for  $x<\alpha$  and  $x>\beta$  and that cohort variation in timing is sufficiently small that  $\alpha-a_m(T)>0$  and  $\beta+a_m(T)<\omega$  for all cohorts.<sup>8</sup>

Figures 1 and 2 show the effects of variation in the model parameters  $g_k(T)$  and  $a_k(T)$  on net marriage rates for men and women. Figure 1 presents the effects of changes in the level parameter  $g_k(T)$  for men ( $k=m$ ) and women ( $k=f$ ) respectively using as a standard the average age-patterns of marriage from the rural Bangladeshi population that will be the subject of our analysis below. The standard male and females schedules are the black dashed and dotted lines, respectively, which correspond to the model schedules evaluated at  $g_m(T)=a_m(T)=0$  and  $g_f(T)=a_f(T)=0$ , respectively. The dark and light grey dashed line represents the model schedule evaluated at  $g_m(T)=.2$  and  $g_m(T)=-.2$ , respectively, and so forth. It is evident from Figure 1 that female age at marriage is earlier and more concentrated than is male age at marriage and that an increase in  $g$  from 0 to .2 results in a proportionate increase in net marriage rates. Figure 2 shows that a change from 0 to 2 in each of the timing parameters shifts the respective marriage schedules to the right by two years.

Because it shifts up the marriage schedule at each age, the cohort-specific parameter  $g_m(T)$  also is related to the overall frequency of marriage by cohort. If the average number of

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<sup>8</sup>This is essentially a technical requirement but is plausible: net marriage rates are essentially zero in early child and at extreme ages, with the latter following in particular because these rates are gross of survival as noted in footnote 2.

marriages per male born in a cohort experiencing the standard marriage schedule is

$$G_{ms} = \int_{\alpha}^{\beta} \phi_{ms}(x) dx \quad (3)$$

then the total number of marriages per male for cohort T,  $G_m(T)$ , is

$$\begin{aligned} G_m(T) &= \int_{\alpha}^{\beta} \phi_m(x, T) dx = (1 + g_m(T)) \int_{\alpha}^{\beta} \phi_{ms}(x - a_m(T)) dx \\ &= (1 + g_m(T)) \int_{\alpha + a_m(T)}^{\beta + a_m(T)} \phi_{ms}(u) du = (1 + g_m(T)) G_{ms} \end{aligned} \quad (4)$$

Thus  $g_m(T)$  measures deviations in the average number of marriages per male cohort member relative to the average number of marriages by men in a cohort experiencing the rates of the standard schedule.

Similarly, if the mean age at marriage of the standard population is  
then the mean age at marriage for cohort T,

$$\begin{aligned} \mu_m(T) &= \int_{\alpha}^{\beta} x \phi_m(x, T) dx / G_m(T) = (1 + g_m(T)) \int_{\alpha}^{\beta} x \phi_{ms}(x - a_m(T)) dx / G_m(T) \\ &= \int_{\alpha + a_m(T)}^{\beta + a_m(T)} (u + a_m(T)) \phi_{ms}(u) du / G_{ms} = \mu_{ms} + a_m(T) \end{aligned} \quad (5)$$

Thus  $a_m(T)$  measures deviations in the average age at marriage for men in cohort T from the average age at marriage for a cohort experiencing the rates given by the standard male schedule.

If variation in the three cohort-specific measures  $c_m(T)$ ,  $a_m(T)$ , and  $g_m(T)$  is small then substantial insight into process of equilibration in the marriage may be obtained by carrying out a linear approximation to the supply of males in these measures, and similarly for women. In particular, relative deviations  $D_m(T) = (M(t) - NG_{ms}) / (NG_{ms})$  in the number males marrying may be written:

$$D_m(t) = \int_0^{\omega} ((g_m(t-x) + c_m(t-x))\phi_{ms}(x) - a_m(t-x)\phi'_{ms}(x))dx \quad (6)$$

which indicates that higher marriage rates and cohort sizes increase the number of males marrying. Assuming, in addition, that  $a_m(T)$  is differentiable, a more interpretable expression may be obtained by integrating by parts

$$D_m(t) = \int_0^{\omega} (g_m(t-x) + c_m(t-x) - a'_m(t-x))\phi_{ms}(x)dx \quad (7)$$

This expression indicates, for example, that decreases in cohort size and overall marriage rates that decrease the number of men marrying at a particular point in time will be offset by decreases in the age at marriage for men. Intuitively, if the age at marriage for men is dropping over time then younger men are entering the marriage market at the same time as their (infinitesimally) older counterparts thus expanding the supply of men to the marriage market at each point in time. While it has been pointed out that changes in the timing of marriage can play an important role in equilibrating the marriage market (Bergstrom and Lam 1993), the simplicity of the relationship has not, to our knowledge been illustrated elsewhere.

A key feature of this relationship is that it is between the *level* of cohort size, on the one hand, and the *change* in the age at marriage: if the number of men marrying in a given year is to be kept constant, then it must be the case that a *temporary* increase in cohort size will result in a *permanent* increase in the age at marriage for men. These relationships may be illustrated by assuming that the standardized marriage schedule is highly concentrated with a peak at some age  $A_{ms}$ .<sup>9</sup> Thus, to first order, the age at marriage for cohort T is  $A_m(T) = A_{ms} + a_m(T)$ , and (7) reduces to

$$D_m(t) = (g_m(t - A_{ms}) + c_m(t - A_{ms}) - a'_m(t - A_{ms})) \quad (8)$$

Assuming further that there is no change in the level of marriage ( $g_m(\cdot) = g_f(\cdot) = 0$ ), we may equate the number of men and women marrying at each time  $t$  and substitute (8), to get

$$a'_m(t^* - \Delta A) - a'_f(t^*) = A'_m(t^* - \Delta A) - A'_f(t^*) = -(c_f(t^*) - c_m(t^* - \Delta A)) \quad (9)$$

where  $\Delta A = A_{ms} - A_{fs}$  and  $t^* = t - A_{fs}$ . Thus if the time derivative of the difference in the ages at marriage (males minus females) equals the negative of the excess fraction of females in corresponding marriage cohorts, the marriage market will remain in equilibrium: a 10% surplus of females may be accommodated through a .1 year decline in the age at marriage difference (e.g, through a .1 year increase in the age at marriage for females assuming  $\Delta A > 0$ ).

Equation (10) also can be used to evaluate the importance of population growth as a

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<sup>9</sup>In particular, we assume that  $\phi_{ms}(x) = \delta(x - A_{ms})$  where  $\delta$  is Dirac's delta. While it may appear that (9) can be obtained directly from (8), this is not precisely true since (8) was derived based on the assumption that the standard schedule  $\phi_{ms}(x)$  is differentiable. The actual derivation is somewhat technical without providing additional insight and is thus omitted. It is available from the author on request.

source of change in the proportion marrying. In the presence of differences in the age at marriage, rapid population growth will tend to create an imbalance between the number of men and women wishing to marry at a particular point in time because, typically, older males from small cohorts are being matched with women from more recent and therefore larger cohorts. Suppose then, that the rate of population growth has been fixed at  $r$  for a long period and that, for simplicity, the male age at marriage is fixed at  $A_{ms}$ . The extent of the imbalance is given to first order by the expression

$$c_m(t - \Delta A) - c_f(t) \approx -r\Delta A \quad (10)$$

Substitution of (9) yields the differential equation.

$$A_f'(t) = r(A_{ms} - A_f(t)) \quad (11)$$

Thus the marriage market can be equilibrated if the female age at marriage is rising at the rate  $r\Delta A$ . Thus, for example, in a population with a growth rate of .025 and an age difference in spouses of 8 years, the female age at marriage would only need to rise by .2 years per year to keep the market in equilibrium. Moreover, as the age-gap narrows the difference in cohort size will fall as well, reducing the magnitude of these effects. Thus, solving (11), yields

$$A_f(t) = A_{ms} - (A_{ms} - A_f(0))\exp(-rt) \quad (12)$$

so that the magnitudes of the age at marriage change required to equilibrate the market will decay exponentially. Thus, for the marriage market to stay in equilibrium in the presence of population growth of this magnitude and an initial age gap of 8 years, one requires only that the age gap fall to 6.23 years over a decade and to 4.85 years over two decades. Arguably these changes are small

enough to not require major changes in the proportion marrying or the operation of the marriage market.<sup>10</sup>

### III. Data

In order to test the ability of this simple model to explain patterns of marriage change we need annual age-specific marriage rates by sex for a reasonably large number of years. While in principle any aggregate data set with accurate reported ages can be used for this purpose, it seems of particular value to examine a country for which the conditions appear to be conducive to the presence of a wide gap in the relative numbers of men and women in the marriage market. Moreover, in order to distinguish marriage-squeeze effects from other changes that might alter the age distribution of marriage it is helpful to have micro-level data so that a structural model of marriage-market allocation may be estimated.

Vital registration data from the records of the Demographic Surveillance System (DSS) of the International Centre for Diarrhoeal Disease Research in Bangladesh (ICDDR) are thus particularly well suited to this analysis. This is an area which has a large gap in age at marriage, as is evident in Figures 1 and 2, and which has undergone substantial population growth over the relevant period yielding a substantial gap between the number of men and women at peak ages at marriage. The population consists of 164,000 people (as of 1974) in 149 spatially contiguous

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<sup>10</sup>The assumption of a roughly even sex-ratio at a given age is relevant here. If the sex-ratio were skewed due to sex selective abortion rather than population growth, a reasonable approximation to the current situation in China (Tuljapurkar et al 1995), then the rate of change in the female age at marriage necessary to equilibrate the marriage market would not diminish with the age gap. Nonetheless the effects would not be substantially different over this time frame: a fixed 20% excess number of males in each cohort would result over two decades in a decrease in the age gap from 8 to 4 rather than 4.85 years. It is worth noting that Tuljapurkar et al's (1995) conclusion that the changes in marriage patterns could not be used to equilibrate the marriage market was based on a comparison of three different fixed marriage patterns and thus did not allow for kind of dynamic adjustment that the analysis presented here suggests is critical.

villages in Matlab Upazilla, which is a rural riverine area. Individual-level, longitudinally collected birth, death and migration registration data are available for all residents of the villages starting in 1966, and all marriage information is available after 1975. Censuses were carried out in 1974 and 1982 and these provide information on household structure, education, and resource availability. All records can be linked at the individual level using permanent individual ID numbers. This not only augments the richness of the data but makes for more accurate age reporting. In particular, ages reported on the marriage certificate have been reconciled with earlier information on age from the censuses.

Marriage records are filed for anyone living in the study area at the time of marriage (whether or not they remain within the area after marriage). These records contain information on education, age and occupation of both spouses. In this study we make use of the marriage records between 1975 and 1990. During this period there were 49,734 unique marriage records reported.<sup>11</sup>

A key issue that has arisen in much of the literature examining the marriage-squeeze and its potential effects on age marriage is the need to define the marriage market appropriately. If the marriage market is defined too broadly, for example, and there is substantial local variation in the relative availability of men and women then marriage-timing may appear to be little affected by variation in cohort size even if, at the local level, this effect is quite important at the local

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<sup>11</sup>Over the 1974-1982 period there are a number of instances in which of marriages are reported twice within a two-year period for the same individuals. These records likely reflect (1) cases in which marriage records were available from both the husband's and the wife's households but were reported in somewhat different time periods (2) some cases in which marriage was never actually consummated (as a result, for example, of the failure of one party to meet their terms of the dowries) and an individual married someone else. In these cases we chose the latter of the two records. There also appears to be a fair amount of divorce (10% of marriages) within two years of the time at marriage, although it is not clear whether these represent consummated marriages or not. In the present study we include these marriages..

level. The presence of detailed vital registration data in a geographically contiguous area provides a particular opportunity to address these issues.

The evidence suggests that it is reasonable to use pre-migration cohort sizes from the study area to characterize the marriage market. First, while there is substantial marriage within villages, the marriage market is clearly broader than the village. In particular, 25.2% of the marriage records villages

In particular, we find that of the 49,734 total marriages over the 16-year period for which at least one spouse was from the study area, 22,754 (45.75%) of husbands and 31,138 (62.6%) of wives were from the study area. The difference between these two area reflects the process of migration. If there were two equal areas, only one of which contains vital registration data, then given patrilocal residence (married couples typically live in the same village if not the same household as the husband's parents (see, e.g., Foster 1992)), the number of women marrying a husband outside of the study area, in the absence of substantial geographic variation in the sex-ratio of marriageable men and women, should equal the number of women from outside marrying inside and thus the fractions of men and women from outside conditional on marriage should be the same. The reason for the discrepancy here is that migration for men typically occurs before marriage, while migration for women occurs at the time of marriage. If the man migrates before marriage and the woman at marriage, only the woman would be recorded as coming from the study area at the time of marriage.

#### IV. Evidence on age adjustment

Table 1 presents results on the mean ages at marriage, based on the marriage-registration data, between 1975 and 1990 as well as the difference in the ages at marriage for spouses. Marriage ages for women in the mid 1970s were extremely low and rose by 3.2 years over the

subsequent 15 year period. As is evident in Figure 1, the mean age at marriage for men is considerably higher<sup>12</sup>: among marriages in 1975, the average age difference between spouses was 10.1 years. Moreover because the male age at marriage only rose by 1.1 years over the study period, the difference in spouses ages declined to 7.7 years in 1990. Indeed, since the male age at marriage barely changed at all until after 1987, the mean age difference for spouses in that year was only 5.92 years.

We turn now to the question of how well these changes in ages at marriage relate to the sizes of the marriage-age population as posited by the standard static model. An overall measure of the relative supply of men and women in each year is constructed using, as peak ages at marriage for men and women, 24-27 and 16-19, respectively. It is evident from the last column of Table 1 that there was a substantial surplus of women during the early part of the study period of almost 75%. This figure reflects the differential in the age at marriage in the population, the rapid population growth, declines in mortality in the early 1960s, and the emigration from the area of young men.<sup>13</sup> More importantly, the extent of the surplus diminishes over time reaching 8.1% in 1987, which also happens to be the year in which the mean age difference is at a minimum.<sup>14</sup>

As is clearly evident from Table 1, the relationship between the relative size of cohorts

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<sup>12</sup>Bergstrom and Bagnoli (1993) provides an explanation of this pervasive phenomenon.

<sup>13</sup>Emigration by age for men and women are not very different in this population: there are approximately the same number of men and women living in the study area at each point in time. Migration plays an important in determining the availability of marriage partners in Matlab because men typically migrate before the peak ages at marriage while women frequently migrate at the time of marriage. As a result, the reported figures effectively overstates the availability of men in the country as a whole.

<sup>14</sup>Note that these quantities do not reflect the changes in the age at marriage (except through the relationship between marriage and migration), because the age ranges are being held fixed.

and the age gap is opposite of what would have been predicted based on the standard static model: a lower excess supply of women is associated with a smaller age gap. The regression line has a slope of .031 and is clearly positive and significant ( $t=5.65$ ).

The resolution is in equation (8). As noted above, the formal demographic model shows that markets can be equilibrated if the *time derivative* in the age at marriage difference (male age minus female age) is positively related to the surplus of men of the appropriate age (or negatively related to the surplus of women). Thus, it is the fact that there is a surplus of women in each year that is potentially related to the rise in the age at marriage for women, not the fact that the extent of the surplus diminished over time. One should expect on this basis that the time derivative of the age at marriage difference should be negative but rising towards zero over time as the surplus of women is reduced somewhat. This prediction is borne out by the fact that the difference in the age at marriage begins to rise just after the surplus female proportion falls below 10%.

To test explicitly the relationship between the surplus proportion of women and the time derivative of the age at marriage difference, it is necessary to translate the period information on marriages into information on the marriage rates of different cohorts, consistent with the formal demographic model discussed above. The difficulty with this approach is that data over a 15 year period provides incomplete information with which to construct estimates of the level and age-profile of marriage of cohorts over the relevant interval. Thus following Foster (1990), we make use of regression methods to extract from the data over the relevant interval information on the timing and level of cohort marriage rates.

We first constructed marriage rates by age and cohort conditional on being in the study population at the age of 13 (the effective minimum marriage age) using the data from the 1975-1990 period. The numerator in each case is the number of marriages by sex of age  $x$  individuals

in that period who are currently resident in the study area and the denominator is the number of individuals of that sex in the corresponding cohort at age 13.<sup>15</sup> These rates were then averaged across cohorts and smoothed using the lowess procedure to obtain a sex-specific standard schedule  $\phi_{ks}(x)$ . The lowess procedure also yields estimates of the first derivative of the standard schedule with respect to age at each age. The smoothed standard rates are plotted, as noted, in Figures 1 and 2. For sex  $k$  we then estimate a linear approximation to equation (2) in the cohort parameters around  $g_k(T)=a_k(T)=0$ :

$$\phi_k(x,T)=(1+g_k(T))\phi_{ks}(x)+a_k(T)\phi'_{ks}(x)+e_{kxT} \quad (13)$$

where  $e_{kxT}$  captures approximation error, by regressing, for each cohort and for the age available for that cohort, the actual cohort rates on the sex-specific smoothed schedule and its first derivative with respect to age. This yields estimates of  $1+g_k(T)$  and  $a_k(T)$ , respectively, as is evident from (13).<sup>16</sup>

The estimates of  $a_k(T)$  and  $g_k(T)$  for men and women of corresponding cohorts (i.e., separated by 8 years, the average age difference between spouses) are presented in Table 2. The range of cohorts for which estimates are provided is determined by the fact that precise estimates of the level and timing of fertility for a particular cohort can only be obtained using this

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<sup>15</sup>For cohorts that reached the age of 13 before the 1974 census we used the sex-specific counts of individuals at the time of that census because previous censuses had somewhat different coverage than the 1974 census and linkage is more difficult. For these cohorts net marriage rates will be somewhat higher and net cohort sizes somewhat lower than would be obtained if cohort sizes at age 13 had been available.

<sup>16</sup>This may be thought of as the first step in iterative non-linear least-squares estimation of (2). Use of a single iteration seemed sensible in that the basic predictions being examined rely themselves on a linear approximation. In practice, subsequent iterations did not substantially alter parameter values in any case.

regression method if the period marriage data are available at ages for which the levels of marriage for that cohort are reasonably high. Because women born before 1960 were over the age of 15 over the entire study period while those born after 1973 were under the age of 17 over the entire period, only cohorts of women born in the years 1960-1973 were included in the analysis. Men were included if they were born 8 years earlier than the corresponding cohort of women (ie., 1952-1965).

The estimates for the timing of marriage for men and women correspond roughly to what was observed for the period data: the cohort mean age at marriage for women rises over the interval at an average rate of .139 years per year, although the largest changes are observed for the cohorts 62-65. Although there is a fair amount of change in the male age at marriage there is little in the way of a trend. In addition the extent of marriage relative to cohort size at age 13 has declined considerably for both men and women, a fact that likely reflects outmigration from the study area rather than a change in the proportion ever marrying or the number of marriages per individual. Although this effect is somewhat stronger for men than women, implying that migration may have worsened the deficit of young men in the study area, the fact that it changes similarly for both limits the importance of this effect with regard to the relative supply of men and women in the marriage market. The final two columns present estimates of the relative surplus of women by cohort<sup>17</sup> and of the time derivative of the difference in age at marriage where the latter is calculated for each cohort T by subtracting the age at marriage difference for cohort T+1 from the age at marriage difference for cohort T:  $(a_m(T+1)-a_f(T+1))-(a_m(T)-a_f(T))$ .

The key question is, what is the relationship is between the time derivative of the age at

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<sup>17</sup>This measure is different from that presented in Table 1 in that it reflects the relative sizes of the corresponding *cohorts* of men and women at the peak ages at marriage.

marriage difference and the relative surplus of women by cohort size? In addition to being presented in Table 2 these measures are graphed in Figure 3. The relationship is extremely strong and negative as suggested: the time derivative of the cohort age at marriage difference in the age at marriage is negatively related to the relative surplus of women (slope=-1.69, t=8.17). The fact that this coefficient exceeds in absolute value the slope that would be expected based on equation (10), -1, may be at least partially attributed to the fact that the level parameter is also changing: a regression that includes differences in marriage probabilities ( $g_m(T)-g_f(T)$ ) yields a coefficient on cohort size of -1.36 that is not significantly different from one. (P=.52).

While it is perhaps not altogether surprising that the data correspond closely to the prediction of the demographic model,<sup>18</sup> the results are nonetheless informative. Faced with a surplus of women of marriageable women for a given schedule the market may respond in several different ways. There may be changes in proportion of individuals marrying, or perhaps increased migration for marriage, or a change in the dispersion of marriage. The fact that the data corresponds so closely to this simple model is at least consistent with the idea that delays in the timing of the marriage are primarily the result of the marriage squeeze. Moreover, it is informative to look at *how* the difference in the age at marriage narrowed: as is evident from equation (10), this narrowing could occur either through earlier marriage by men or through later marriage by women. While these results cannot tell us exactly why later ages at marriage were primarily responsible, they do show that the response of the women was dominant.

## V. Behavioral Model

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<sup>18</sup>The fit suggests that key assumptions used to derive (10) are not unreasonable. Of particular importance are the assumption that the market clears at the local level, that cohort data can be well approximated by a two parameter model, and that variation in parameters is sufficiently small that a linear approximation is appropriate. .

To gain some insight into the behaviors that underlie the patterns observed in Tables 1 and 2, it is helpful to know how potential spouses are valued in the marriage market. In effect what is needed is a price that reflects how different attributes are rewarded at each point in time. One approach to this problem is to make use of information on dowries as in Rao (1993). Unfortunately, detailed information on dowries is not available in these data. In any case, given the complexity of marital transactions it is unclear whether dowries fully measure differences in the value of potential spouses: intrahousehold allocations and post-marriage interhousehold transfers (Rosenzweig and Stark 1989) may substitute for the payment of dowries. Thus we need a method for constructing implicit measure of the value of different potential spouses.

In order to construct such an estimate we consider a simple economic model. In particular, we make use of transferable utility with private consumption goods for each spouse and a public (within marriage) good which is an adaptation of the model used by Bergstrom and Lam (1993) to simulate age at marriage change using aggregate Swedish data.<sup>19</sup> As noted by Bergstrom and Lam (1993), this model exhibits two convenient properties: (i) the level of public good consumption within each marriage is not influenced by distributional considerations and (ii) any pareto efficient allocation in the marriage market maximizes the sum of total utility.

In particular we assume that the utility of a married male  $i$  may be written

$$u_{mi}(c, h) = c - \frac{1}{2}(h_m - h)^2 \quad (14)$$

where  $c$  is the private consumption good,  $h$  is the consumption of the public good, and  $h_m$  is a

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<sup>19</sup>In their approach the public good is the time of marriage which is directly observable. In ours the public good is not directly observable but can be modeled, as shown, below as a function of partner-specific observables.

parameter that represents the unconstrained optimum consumption of the public good.<sup>20</sup> Utility is increasing in private consumption and decreasing in the distance between the actual and ideal level of consumption of the public good. A similar utility functions holds for females with the subscript m replaced with an f.

We assume further that the public good is only produced at home and that the cost of producing the good is linear and decreasing in the human capital (broadly construed) of the husband  $r_{mi}$  and wife  $r_{fi}$  respectively. For concreteness it is helpful to think of the public good as being child human capital, in which case what is being assumed here is that the cost of producing a given level of human capital for a child will be lower if that child's parents have higher levels of human capital themselves. Thus the budget constraint that arises when male i marries female j is:

$$c_i + c_j + (\gamma_0 - \gamma_m r_{mi} - \gamma_f r_{fi}) h_{ij} = y_{mi} + y_{fi} \quad (15)$$

where  $c_i$  and  $c_j$  are the parents' respective level of private consumption,  $h_{ij}$  is the level of public good selected,  $y_{mi}$  and  $y_{fi}$  are the parents' respective levels of earnings, and  $\gamma_0 - \gamma_m r_{mi} - \gamma_f r_{fi}$  is the price of child services.

Under these conditions any efficient allocation given the attributes of i and j has the property that the level of public good that will be chosen is (Bergstrom 1997):

$$h_{ij} = (h_m + h_f - (\gamma_0 - \gamma_m r_{mi} - \gamma_f r_{fi}))/2 \quad (16)$$

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<sup>20</sup>We assume, for simplicity, that the ideal level of child human capital varies only by sex. Relaxing this assumption would change the interpretation of the estimated coefficients but would not alter the proposed estimation or the resulting implications for the relative value of different criteria in the marriage market.

and that the sum of utilities of the man and woman in the couple is

$$Z_{ij} = z_m(r_{mi}y_{mi}) + z_f(r_{fi}y_{fi}) - \frac{1}{2}\gamma_m\gamma_f r_{mi}r_{fi} + \epsilon_{ij} \quad (17)$$

where  $z_m()$  and  $z_f()$  are quadratic functions and  $\epsilon_{ij}$  is an added match-specific shock that may be thought of as “chemistry”. The third term in equation (17) says simply that, *ceteris paribus*, joint utility will be higher if a high human capital male is married to a high human capital female.

As noted, a key implication of the transferable utility function is that, under these conditions, the marriage market will operate in a way that maximizes the total amount of marital utility given the individuals who choose to marry at a particular date (Bergstrom 1997). The basic idea is that distributions within marriage do not affect which marriages take place because the distributional issues can be dealt with through a reallocation of private consumption. The fact that total welfare is maximized overall implies that given any two couples within this marriage market that the total welfare of these four individuals could not have been improved by a swap of their respective spouses. In particular, for males  $i$  and  $k$  and females  $j$  and  $l$ , with  $i$  married to  $j$  and  $k$  married to  $l$ ,

$$Z_{ij} + Z_{kl} - Z_{il} - Z_{jk} \geq 0 \quad (18)$$

which requires

$$\epsilon_{ij} + \epsilon_{kl} - \epsilon_{ik} - \epsilon_{jl} \leq (r_{mi} - r_{mj})(r_{fk} - r_{fl}) \quad (19)$$

Note that the functions  $z_m()$  and  $z_f()$  in (17) drop out of (19) because they appear linearly in  $Z$  and depend only on the characteristics of individuals, not of couples. Also, the fact that the male and female attribute equations appear in this expression in differences implies that time-specific

unobservables that are shared by all those marrying in a particular year (including those related to the selection into marriage in that period) will be differenced out.

We assume that the parental human capital variables are linear functions of observables individual characteristics that may influence parental productivity in the provision of child human capital with the coefficients differing by sex:  $r_{mi} = \beta'_m x_{mi}$  and  $r_{fi} = \beta'_f x_{fi}$ , for men and women, respectively, where  $\beta_m$  and  $\beta_f$  are coefficient vectors and  $x_{mi}$  and  $x_{fi}$  are vectors of characteristics. We also assume that the match specific shock  $\epsilon_{ij}$  is distributed such that the sum  $\epsilon_{ij} + \epsilon_{kl}$  has an extreme value distribution. Let  $I_{ijkl}$  be an indicator which takes the value 1 if  $i$  marries  $j$  and  $k$  marries  $l$ . Then

$$Prob(I_{ijkl} = 1 | I_{ijkl} + I_{ijk} = 1) = 1 / (1 + e^{-(\beta'_m(x_i - x_k))(\beta'_f(x_j - x_l))}) \quad (20)$$

which suggests that the parameter estimates may be obtained using a non-linear conditional logit procedure.

The vector of characteristics included in  $x$  are, for each individual, the number of items owned by the household,<sup>21</sup> education, head's education, age and age squared. A normalization is also necessary: doubling the coefficients on  $h_m$  and halving the coefficients on  $h_f$  would not change the likelihood. Thus, we set the coefficient on the male items code to one. Under the assumption that an increase in the number of items owned by the male's household makes him more attractive in the marriage market, we may thus interpret the other estimated coefficients as

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<sup>21</sup>The number of items code has generally been found to be a relatively effective measure of socio-economic conditions in this population. The code is the sum of indicator variables indicating whether the household owns a radio, a watch, a hurricane lamp, a traditional quilt, and whether it receives remittances. This measure thus has a maximum of 5 and a mean in 1974 of 2.8. Unfortunately information on land ownership was not collected in the 1974 census.

measures of the extent to which a given attribute raises an individual's attractiveness in the marriage market.<sup>22</sup>

The resulting coefficient estimates are presented in Table 3. In each column the coefficients on the wife's attributes are presented (i.e., the estimates of  $\beta_f$ ), followed by the coefficients for the husband's attributes,  $\beta_m$ ). At the bottom of the table, the estimates of the most preferred age at marriage for each sex (computed based on the quadratic age effects) are also presented.

The first column pools the sample over the 1975-1990 period. The estimates appear quite sensible: women from households with more items owned, with a more educated head, and who are themselves more educated all are apparently more attractive in the marriage market than those from poorer households with lower levels of schooling. Education proved to be quite important: the value of having a secondary educated head is roughly comparable to the value of having 2.3 additional items. The education of the woman appears to be at least as important as the education of her father: a woman with no schooling and a father with more than primary schooling would be as preferred as one with some schooling and an uneducated father.<sup>23</sup>

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<sup>22</sup>Note that in the spirit of our model, "attractiveness in the marriage market" refers to the extent of human capital of an individual, and operates only through the effects of human capital on the cost of the provision of the public good. There is nothing in the model that would imply that there should be assortative mating according to, for example, income  $y(x)$  other than to the extent that  $h(x)$  and  $y(x)$  are closely associated with each other (as we might expect to be the case in practice). Indeed, given  $h_f(x)$ , a potential groom will be indifferent to marrying women with different  $y_f(x)$  because the extra resources of a higher income wife will be transferred away in the form of a dowry or through greater consumption by the wife.

<sup>23</sup>Preliminary estimates that distinguished between primary and less-than-primary schooling for the women indicated that having more than 5 years of schooling yielded a very large effect; however, given the fact that very few women in the early period were educated beyond the primary level, these estimates proved to be quite unstable when stratified by period and thus we have chosen to condition only on having some education.

Evidently the degree of preference for women of a particular age is not very strong: the coefficients on age indicate a deviation of plus or minus five from the preferred age at marriage is equivalent to only a .3 drop in the number of items.

The results for husbands are broadly similar. In this case it is clear that the characteristics of the spouse are much more important than the characteristics of head: a spouse with some primary education from a head with no education is preferred to a spouse with no education from a household where the head has completed primary education. Relative to the number of items owned in the household (recall that the coefficient on items for the husband has been normalized to one) the effects of some primary education for the husband (2.01) are slightly less than the effects observed for a woman with some education (2.16). Finally, the costs of not marrying at a particular age are roughly the same for men as for women: a five year deviation from the preferred age at marriage is equivalent to a .43 drop in the number of items.

Given the rapid change in the age at marriage over time, it is instructive to ask whether there was a change in how women and men were evaluated in the marriage market over time. It has been suggested (Caldwell *et al.* 1983, Lindenbaum 1981) that a rise in the value of husband's education from the perspective of potential brides (or their families) may have played an important role in the rise in education as well as a shift from bride prices to dowries. Also, consistent with the spirit of our model, an increased interest in providing children with schooling in the area may have increased the desirability of having a mature and educated wife compared with the value of having a wife from a good household.<sup>24</sup>

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<sup>24</sup>Foster and Roy (1999) discuss the decreases in fertility and increases in educational investment in the Matlab population over the relevant period. Behrman, Foster, Rosenzweig and Vashistha (1999) and Foster and Rosenzweig (2000) show that increases in the returns to male human capital increase the demand for educated mothers as an input in the production of high human capital children. Thus decreases in fertility may, through a quality-quantity tradeoff, increase the

A comparison between columns two and three indicates that any changes were quite moderate. There is a small rise in the extent to which the education of women was valued, particularly compared with the extent to which the education of the head in her household was valued, which is consistent with the idea that the characteristics of the women became more valued than the characteristics of her household. There is little change in the relative value of education of husbands, however.

The most dramatic changes occur with respect to the extent to which individuals of different ages were valued. In particular, the preferred age at marriage for men rose from 24.3 to 29.6 years while the preferred age at marriage for women rose from 15.7 to 19.6 years. Thus, the desired age at marriage shifted sharply upward for members of both sexes. Given the scarcity of the number of older unmarried men 25-30, relative to unmarried women 16-20, this shift in preferences likely had a relatively pronounced effect on the opportunities open to older, unmarried women, thus allowing age at marriage to equilibrate the market. Put another way, if preferences had remained the same, then women who delayed marriage in the 1970s because of the adverse circumstances vis a vis the availability of potential spouses would at least have had to accept less attractive marriage partners if they had married at all. The fact that the preferred age at marriage shifted implies that these women did reasonably well in the marriage market.<sup>25</sup>

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importance given to schooling in the marriage market. To the extent that increases in age may improve maternal quality this same mechanism might also increase the optimal age of brides.

<sup>25</sup>It is worth pointing out that in this model dowries play no role in the determination of who marries whom: they simply reflect the allocations of resources given marriage. In a less restrictive model, older women may appear to be more attractive marriage partners because they come with larger dowries, which may themselves reflect a greater willingness of parents to pay to marry off their older daughters out of concern that the daughters may soon become too old to marry. But this seems at odds with the fact that the relative surplus of women is falling as the most attractive age at marriage rises.

## VII. Conclusion

In this paper we have undertaken an examination of a rapid rise in the female age at marriage in a rural area of Bangladesh. The results suggests that the primary mechanism underlying this change in the age at marriage in that population was the relative scarcity of men of marriageable age in the population rather than any fundamental change in economic or social conditions. The model also explains why the increase in the age at marriage of women observed in Matlab in the 1975-90 period corresponded to a period in which the magnitude of the surplus of women actually decreased, a result that seems to be at odds with most of the discussion of the relationship between cohort size and the age at marriage gap in the literature.

In the course of our analysis we have used a simple demographic model to show that, consistent with what was observed in the study population, the marriage market can easily accommodate changes in cohort size. We have also obtained estimates of the attractiveness of different potential spouses in the context of an economic model of marriage that emphasizes the role played by the attributes of one's partner in the provision of a household-level public good. The estimates suggest that individuals in this population do not have especially strong preferences with respect to the age of their spouse at marriage and that the preferred age of a spouse at marriage increased for both sexes over the relevant period. To the extent that these results generalize to south Asia as whole, one may question the role of a marriage squeeze in the transition from bride price to dowry in parts of south Asia. If there are not strong preferences with respect to age at marriage and if preferences over time have shifted to higher desired ages then marriage-market payments might be expected to be relatively unresponsive to the relative numbers of men and women at peak marriage ages. More generally, the results suggest that concerns expressed in the literature over the adverse consequences of imbalances in the sex ratio

at peak ages at marriage and the inability of the marriage market to adjust to these conditions through changes in the age at marriage have been overstated.

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Table 1  
 Period Mean Ages at Marriage and Numbers  
 of Individual at Peak Marriage Ages

Year	Mean Ages at Marriage			Number at Peak Marriage Ages		
	Women	Men	Difference	Males (24-27)	Female (16-19)	Ratio of Women to Men
75	16.8	26.9	10.1	3504	6081	1.735
76	16.9	25.9	9.02	3555	6325	1.779
77	17.3	26.2	8.87	3766	6753	1.793
78	17.4	26	8.65	4064	6753	1.662
79	17.6	26.1	8.42	4122	6776	1.644
80	17.8	26.4	8.51	4418	7138	1.616
81	17.9	26.1	8.12	5167	7177	1.389
82	18.3	26.1	7.79	5681	7903	1.391
83	19.2	27.4	8.2	6264	8318	1.328
84	19.4	27.3	7.94	6659	7542	1.133
85	19.6	27.1	7.46	6549	7308	1.116
86	20.2	26.4	6.08	6738	7062	1.048
87	20.2	26.6	5.92	6532	7061	1.081
88	20	27.7	7.71	6373	7145	1.121
89	20.1	27.9	7.8	6103	7358	1.206
90	20.3	28	7.7	6041	7292	1.207

Table 2

Estimates of Cohort Timing and Level Parameters for Male and Female Births Cohorts Marrying in Approximately the Same Year							
Birth Cohorts Separated by 8 years		Level Parameter a(T)		Level Parameter g(T)		Ratio of Women to Men	Change in Age at Marriage Difference $\Delta(a_m(T) - a_f(T))$
		Women	Men	Women	Men		
60	52	-0.853	0.841	0.304	0.621	1.754	-0.938
61	53	-1.03	-0.274	0.263	0.661	1.745	-0.892
62	54	-0.954	-1.09	0.216	0.653	1.698	-0.619
63	55	-0.675	-1.43	0.209	0.576	1.612	-0.431
64	56	-0.354	-1.54	0.161	0.541	1.508	-0.489
65	57	-0.025	-1.7	0.089	0.416	1.391	-0.422
66	58	0.227	-1.87	0.014	0.259	1.288	-0.04
67	59	0.357	-1.78	0.011	0.039	1.198	0.352
68	60	0.365	-1.42	-0.011	-0.084	1.134	0.532
69	61	0.323	-0.93	-0.037	-0.151	1.103	0.395
70	62	0.452	-0.406	-0.1	-0.168	1.091	0.136
71	63	0.639	-0.083	-0.16	-0.197	1.111	0.009
72	64	0.679	-0.034	-0.212	-0.244	1.106	0.07
73	65	0.493	-0.15	-0.258	-0.31	1.087	--
Average Annual Change (t-ratio)		.139 (9.27)	.022 (.387)	-.043 (30.7)	-.088 (13.67)	-.060 (10.48)	--

Table 3  
Nonlinear Logit Estimates of Marriage Market Valuations<sup>a</sup>

Wife	1975-1990	1975-1982	1983-1990
Items Owned	.074 (6.91) <sup>b</sup>	.07 (5.05)	.080 (4.68)
Educ:			
1-5 yrs	.160 (5.63)	.147 (4.92)	.178 (4.36)
Heads Educ:			
1-5 yrs	.104 (5.16)	.116 (4.18)	-.036 (.584)
5+.	.170 (5.63)	.171 (4.13)	.115 (3.58)
Age (x10 <sup>-1</sup> )	.081 (1.20)	.117 (1.39)	.070 (.769)
Age (x10 <sup>-2</sup> )	-.099 (2.48)	-.152 (2.53)	-.036 (.584)
Husband			
Educ:			
1-5 yrs	2.01 (5.17)	1.94 (3.62)	2.02 (3.61)
5+	5.52 (6.56)	5.42 (4.83)	5.52 (4.38)
Heads Educ:			
1-5 yrs	.519 (1.76)	.421 (1.06)	.591 (1.36)
5+.	1.72 (3.42)	5.42 (4.83)	1.08 (1.62)
Age (x10 <sup>-1</sup> )	4.00 (3.55)	5.44 (3.03)	3.31 (2.18)
Age <sup>2</sup> (x10 <sup>-2</sup> )	-1.62 (4.04)	-2.41 (3.43)	-.997 (1.85)
Implicit preferred age:			
Wife	17.09	16.85	22.72
Husband	0.65	1.71	29.60
Average log-likelihood (N)	-.610 (3914)	-.621 (1832)	-.596 (2082)

<sup>a</sup>Coefficient on the number of items in husband's household normalized to one.

<sup>b</sup>Asymptotic t-ratios in parentheses

Figure 3  
Effects on Ratio of Women to Men on Time  
Derivative of Age at Marriage Difference

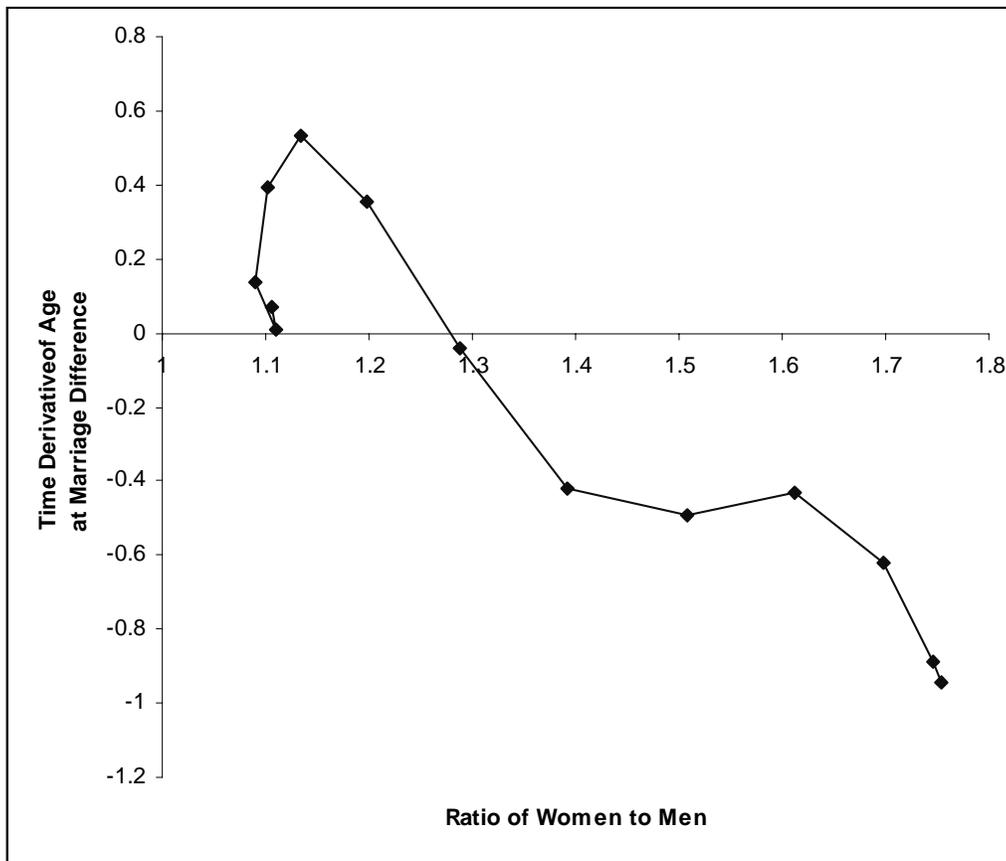


Figure 1  
Effects of changes in the timing parameter on predicted marriage rates

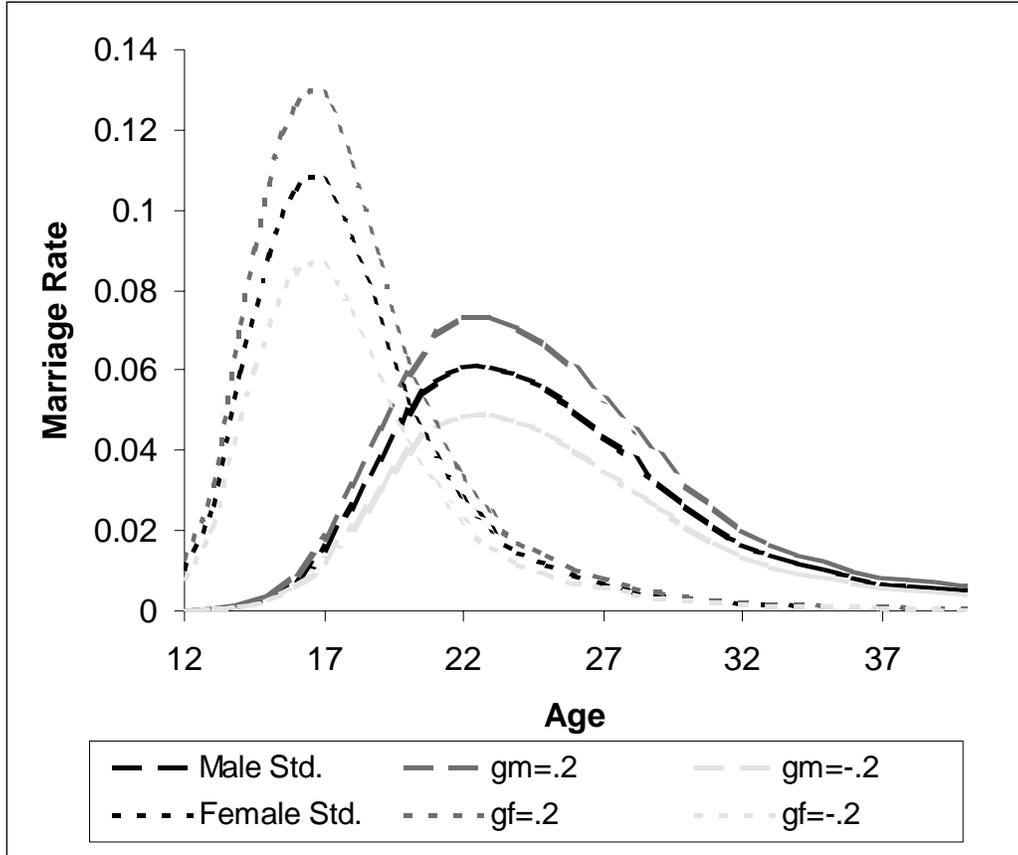


Figure 2  
Effects of changes in the timing parameter on predicted marriage rates

