Male Earnings Inequality and Female Marital Outcomes: Evidence from India

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Abstract

We provide the first evidence, for a developing country, of the causal effect of widening income inequality on the age of marriage of women. We extend the logic of the job search model. In this case, greater income inequality makes high-earning men relatively more attractive candidates than before, and increases the willingness of a woman to wait for such candidates, thus delaying her marriage. We utilize data from India to show that this mechanism does indeed have appear to have increased the average age at marriage. Further, it seems that women are using the extended search duration to acquire more education. We are able exploit the richness of the data to rule out a number of alternative explanations for our results. The results are not due to (i) Men searching longer in the marriage market in response to greater female earnings inequality, (ii) Regional or caste-base social norms, (iii) Men searching longer in the labor market (reducing the gender ratio in the marriage market), (iv) Earnings dispersion proxying for educational premia, and hence encouraging women to stay in school longer, or (v) Women's families needing more time to afford greater dowries.

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1 Introduction

The phenomenon of women marrying young is widespread and is of acute significance in developing countries. In addition to its role in the fertility transition, studies have found that a women's age at marriage correlates strongly with a number of health outcomes. Early marriage has been associated with low contraceptive use, miscarriages, multiple unwanted pregnancies, domestic violence, depression, and even increased HIV risk (Bruce, 2003; Clark, 2004; Nour, 2006; Raj et al., 2009; Santhya et al., 2010). Data from various Demographic and Health Surveys (DHS) conducted between 2003 and 2005 reveal that as many as 27.1% of women in India between the ages of 15 and 19 are married (the legal minimum age is 18 years). The issue is more severe in other developing countries, with corresponding figures of 46.1% in Bangladesh, 42% in Chad, 32.9% in Malawi, 50.4% in Mali, 38.2% in Mozambique and 31.7% in Nigeria.

However, the age at which women marry has been gradually increasing, although there is some heterogeneity in these trends across countries in this regard (Singh and Samara, 1996; Harwood-Lejeune, 2000; Choe et al., 2002; Heaton et al., 2002; Rashad and Osman, 2003; Westoff, 2003; Jensen and Thornton, 2003). In India, from 1980-2005, women's age at marriage increased by 4.3 years (Figure 1). The trend towards later marriage has been hypothesized to reflect factors such as access to education. However, Figure 2 indicates that in the case of India, more education cannot fully explain the increase of the marriage age. From 1980-2005, married women acquired an additional 1.5 years of education, which explains less than half the increase in the marriage age over the same period. Other explanations for delayed marriage include improved labor market opportunities, demographic change, urbanization, and exposure to mass media (see, for example, Jensen and Thornton, 2003; Mensch et al., 2005).

A less obvious mechanism, and one that has not been evaluated in the developing country context, is that increasing income inequality may delay marriage decisions. This mechanism may be understood in the context of a search model in the marriage market as in Loughran (2000 & 2002). In this model, a woman who is on the marriage market samples marriage offers from prospective spouses, evaluating them in terms of their income. In deciding whether to accept a particular offer, she must decide whether the benefits from waiting for a better offer outweigh the costs of staying unmarried during this period. Widening income inequality makes high-earning men relatively more attractive candidates than before, and increases the willingness of the woman to wait for offers from such candidates, thereby delaying her marriage. Loughran (2002) and Gould and Paserman (2003) provide evidence that this mechanism has significantly reduced marriage rates in the United States over the last few decades.

Utilizing nationally representative household survey data from India, we provide the first evidence for such a phenomenon in the context of a developing country. The role of income inequality in delaying marriage is particularly relevant in the Indian setting because the turn towards economic liberalization taken during the 1990s appears to have increased income and consumption inequality, both over time as well as geographically (Banerjee and Piketty, 2005; Deaton and Dreze, 2002; Sundaram and Tendulkar, 2003a & 2003b; Sen and Himanshu, 2005; Pal and Ghosh, 2007). The available evidence indicates that this growing inequality has been largely driven by an increase in incomes at the upper tails of the income distribution, while incomes of other groups have stagnated. Banerjee and Piketty (2005) use income tax data to show that incomes of the top 1% in India increased by 50% in the 1990s, and that their share of total income nearly doubled.

To evaluate the link between earnings inequality and marriage in India, we relate (age-specific) marriage rates of women of marriageable age to male earnings inequality in the relevant marriage market, which we define as all out-of-school unmarried males of marriageable age belonging to the same caste (social group) and residing in the same state. This definition of the marriage market takes into account the fact that inter-caste marriages in India are rare, and that ethno-linguistic differences between states limit the geographical boundaries of the marriage market. This natural segmentation in India provides substantial variation in the marriage market experiences of women, and we are able to utilize this variation to establish the effect of earnings inequality on marriage while controlling for spatial and inter-caste socio-economic characteristics that may influence marriage decisions.

Consistent with the predictions of the marital search model, we find that women in marriage markets characterized by higher male earnings inequality have significantly lower (age-specific) rates of marriage. Men, on the other hand, do not delay their marriage as a response to higher female earnings inequality. While Loughran (2002) and Gould and Paserman (2003) examine the effect of income inequality on marriage rates, we are able to go a step further and establish the extent of time marriage is postponed. We can do so because our data contain a subsample of married women for whom we know the age at marriage. The results indicate that a doubling of the earnings of men at the 90th percentile relative to median male earnings reduces the (age-specific) propensity of women to get married, resulting in an average delay of marriage by nearly 2 years.

Marriage in India (and many other developing countries) is strongly linked with education. In the majority of cases, schooling is discontinued immediately upon cohabitation, if not much earlier. Hence, an important benefit of postponing marriage in the developing country setting is that it can increase female educational attainment (Brien and Lillard, 1994; Lindstrom and Brambila Paz, 2001; Fields and Ambrus, 2008; Maertens, 2013). We find this to be the case: women in marriage markets with greater male earnings inequality also obtain more education. This additional education is concentrated at the point in the educational trajectory at which most women find themselves when they are on the marriage market, i.e., greater male earnings inequality results in increased rates of matriculation and high-school graduation among women, but does not correlate either with the probability of receiving zero education or the probability of completing primary school. We also show that greater earnings inequality does not affect the educational attainment or school enrollment of children who are too young to be on the marriage market. These results confirm that male earnings inequality impacts female educational attainment only at the time of marital search and not before.

We conclude that there is a strong link between male earnings inequality and female age at marriage, and that this relationship plausibly reflects the causal effect of the former on marital search duration. In turn, the increase in search duration appears to improve educational attainment, resulting in significantly higher rates of matriculation and high-school completion. The results in this paper suggest the strong possibility that widening income inequality in India may have had some salutary effects working through the marriage market.

The paper proceeds as follows. Section 2 outlines a model of marital search to motivate the empirical analysis. Section 3 describes the data and the variables used in the analysis. Sections 4 and 5 present the empirical strategy and the results. Section 6 explores alternative hypotheses and Section 7 concludes.

2 A model of marriage market search

This is a partial equilibrium model, that focuses attention on how changes in the characteristics of the male earnings distribution affect the marital search duration of women. This section presents a model that is used to derive predictions about how the duration of a woman's marriage market search is related to the mean, variance and upper-tail characteristics of the male earnings distribution. In substance, the model is no different from a standard labor market search model. For simplicity of exposition, therefore, we present a version of the model that embodies some functional form assumptions, while referring readers to Mortensen (1986) for the derivation of the results in the more general case.

We model the marriage market search as an extension of a labor market search model (see among others, McCall (1970)), in which an infinitely-lived woman sequentially samples marriage offers from a known distribution of offers, before finally accepting an offer. As in Loughran (2002), we assume that women prefer high-earning males over low-earnings males and that this is the only dimension of groom heterogeneity, so that the male earnings distribution characterizes the distribution of offers. However, the empirical analysis allows the marriage market to be structured horizontally along another important dimension: caste.

The asymmetric nature of the model fits into the Indian context. Traditionally, when a female child reached the age of menarche, the child was taken out of school and arrangements for her marriage were made (Caldwell et al., 1983). Nowadays, while parents in India have more leeway, they do become anxious to marry off their daughter once she has reached menarche, partly to avert any unwanted pregnancies (see also Srinivas, 1984). Once a daughter is considered of marriageable age, inquiries are sent across the (caste-based) social network. Suitable candidates present themselves, and out of this set, a selection is made. Maertens (2013), using data from three villages in India, finds that families meet up to 5 candidates in this process. While the prospective spouse is usually only met on the day of the wedding (65% of the women in our sample first met their husband at the wedding), 42% of the brides do have a say in the selection process. From the point of view of the men, Maertens (2013) finds that education and settling down (e.g., getting a regular job or establishing a household) receives preference over marriage; so usually only out-of-school men could be considered as potential candidates, and only men with certain income and wealth prospects would be considered desirable matches.

2.1 Set-up

Suppose F is the cumulative distribution function corresponding to the male earnings distribution, and let f denote the associated density function. For simplicity, we assume that the woman receives exactly one offer in each period and that this offer is an independent and random draw from F. We denote the earnings associated with this offer by x. The woman may either accept or reject this proposal. If she chooses to reject it, she will have to wait one more period, during which she will receive a utility of c, which is the per-period utility from remaining single, and may be negative for some women. Recognizing the iterative nature of the problem we follow a traditional dynamic programming framework.

Denote by $V_{accept}(x)$ the lifetime value obtained by accepting offer x. Assuming that the woman is risk-neutral, we can write:

$$V_{accept}(x) = \frac{x}{1-\beta} \tag{1}$$

where β is the discount rate.

Denote by V_{reject} the value of rejecting offer x. Because of the recursive nature of the problem, we can write:

$$V_{reject} = c + \beta \mathbb{E}\left[\max\left\{\frac{x}{1-\beta}, V_{reject}\right\}\right] \tag{2}$$

The form of the solution to this optimization problem can be shown to be the following: there is a reservation earnings level, denoted by R, above which the woman will accept any offer, and below which she will reject any offer and instead choose to wait another period. We can therefore re-write (2) as:

$$\begin{aligned} V_{reject} &= c + \beta\left[\int_R^\infty \frac{x}{1-\beta} dF(x) + \int_{-\infty}^R V_{reject}\, dF(x)\right] \\ &= c + \frac{\beta}{1-\beta}\int_R^\infty x dF(x) + \beta F(R) V_{reject} \end{aligned} \tag{3}$$

We also note that if the offer exactly equals the reservation level, R, the woman will be indifferent between accepting and rejecting, i.e. we must have:

$$V_{reject} = V_{accept}(R) = \frac{R}{1-\beta} \tag{4}$$

Combining (3) and (4), we can write, after some algebra:

$$[r + 1 - F(R)]R = rc + \int_{R}^{\infty} x dF(x) \tag{5}$$

where $r = \frac{1-\beta}{\beta}$.

This equation may be solved to obtain the reservation earnings level, R. Associated with R, one can define the per-period probability of "escape" from the marriage market, i.e., the probability of getting married, q:

$$q = P(x > R) = 1 - F(R) \tag{6}$$

Lower values of q imply greater search duration (in expectation), or equivalently, higher age at marriage. The expected number of periods of search before an offer is accepted can be shown to be $1/q$.[1]

2.2 Comparative Statics

We now present the key comparative statics of interest. We note that the outcome of interest is the expected search duration, which is inversely related to q. The comparative static results presented here are derived by assuming that F is $N(0, 1)$. This allows us to simplify the algebra. It can be verified however that the results obtained here are fully general, and are not dependent on the normality assumption (see Mortensen, 1986).

First, we consider the effect of changes in the per-period utility (c) on the per-period escape probability (q), while holding fixed the earnings distribution (F). Note that in this case it is sufficient to know the effect on R to infer the effect on q (this follows from Equation (6) above). For different levels of c, the reservation earnings will be different. Applying the Implicit Function Theorem to equation (5) and using (6), one can show that higher levels of c correspond to higher

[1] Denote search duration by T. Then, $\mathbb{E}(T) = 1 * P(T = 1) + 2 * P(T = 2) + 3 * P(T = 3)... = q[1 + (2(1-q) + 3(1-q)^2 + ...]$
Denote $S = 1 + 2x + 3x^2 + 4x^3 + ...$with $x = 1 - q$. We can show that $S = 1/q^2$ and hence $\mathbb{E}(T) = 1/q$.

values of R and hence lower values of q:

$$\frac{dR}{dc} = \frac{r}{r+q} > 0 \tag{7}$$

This result says that women who obtain little utility (or great disutility) from remaining single are less likely to turn down offers.

Next, we examine how changes in the earnings distribution, F, affect search duration. A complication in this case is that when F changes, the effect on search duration cannot automatically be inferred from the change in R.

We begin by considering the effect of a marginal rightward shift in the earnings distribution, while holding constant the variance, i.e. the mean of the distribution increases. Set $F(x) = \Phi(\frac{x-\mu}{\sigma})$ where Φ denotes the standard normal distribution function. Again, applying the Implicit Function Theorem to equation (5) utilizing (6) yields (see also Appendix):

$$\frac{dR}{d\mu} = \frac{q}{r+q} < 1 \tag{8}$$

And, using (6):

$$\frac{dq}{d\mu} = \left(1 - \frac{dR}{d\mu}\right)\phi(R) > 0 \tag{9}$$

where ϕ denotes the standard normal density function. Thus, a rightward shift in the offer distribution will increase the reservation level, R. However, R increases less than one-for-one with mean μ, and in turn this implies that the per-period escape probability increases, and so the expected age at marriage must fall. Because r is likely to be quite small relative to q, we expect this effect to be small.

Next, we consider a mean-preserving spread in the offer distribution. Because of our assumption of normality, this is equivalent to an increase in the standard deviation of the offer distribution, while holding constant the mean. It is important to note that except in this special case, the equivalence does not hold. Mortensen (1986) provides the comparative static result in the general case of a mean-preserving spread starting from an arbitrary distribution.

Applying the Implicit Function Theorem to equation (5) using (6) yields (see also Appendix):

$$\frac{dR}{d\sigma} = \frac{\phi(R)}{r+q} > 0 \tag{10}$$

And, using (6):

$$\frac{dq}{d\sigma} = \left[R - \frac{dR}{d\sigma}\right]\phi(R) \tag{11}$$

The increase in earnings dispersion, σ, unambiguously increases the reservation level, R. However, this time, the direction of the effect on q is ambiguous, depending on the term in square brackets in Equation (11).

To understand these results, note first that for any initial value of R, the expected earnings of individuals above R increases with a mean-preserving spread of F. This implies that the value of rejecting an offer equal to R has increased, which in turn will cause the woman to revise R upwards (because initially, the value of accepting an offer equal to R was equal to the value of rejecting it). Thus the reservation level unambiguously rises for all women, regardless of their initial value of R.

Next, note that the mean-preserving spread increased the probability mass to the right of R (i.e. the escape probability, $P(x > R)$) in the case of women for whom R was above the median, while decreasing $P(x > R)$ in the case of women for whom R was below the median. For the latter group, as R rises in response to the mean-preserving spread, $P(x > R)$ falls further, whereas for initial reservation values above the median, the net effect of the increase in R on $P(x > R)$ is ambiguous, and depends on how large the increase in R is. For this reason, the mean-preserving spread has a theoretically ambiguous effect on search duration for women whose initial reservation levels were above the median.

Nevertheless, it is likely that for most values of R, the net effect on search duration is positive, i.e. q falls (see also Burdett and Ondrich (1985) for discussion on this point). Figure 3 illustrates this in the context of our particular functional form assumptions, by graphing $R - \frac{dR}{d\sigma}$ against different values of R (where we have assumed a discount rate of $\beta = 0.95$). The figure also graphs the standard normal cumulative distribution function. The figure shows that $R - \frac{dR}{d\sigma}$ is indeed negative over a substantial range of values of R, accounting for a large proportion of the probability mass. It is only in the far right tail of the distribution, high values of R, that $R - \frac{dR}{d\sigma}$ (and hence

$\frac{dq}{d\sigma}$) turns positive. Nonetheless, the question of whether increasing earnings dispersion increases search duration and the age at marriage is ultimately an empirical one, and will be tested in the data.

We have so far considered a mean-preserving spread as an example of rising inequality. However, we are unlikely to observe mean-preserving spreads in the data. The available evidence also indicates that for a number of English-speaking countries as well as India and China, recent increases in income inequality are largely being driven by increases in income for people in the upper tails of the income distribution (Atkinson, Piketty, and Saez, 2011). In the empirical analysis, we will accordingly separate out the impact of increases in upper-tail incomes (relative to median incomes) from the impact of reductions in lower-tail incomes.

While analytical proofs are difficult, it is possible to intuitively predict the direction of impact in these cases. Consider an increase in income at the 90th percentile, relative to the median income. For women with reservation levels below the 90th percentile (this includes the majority of women), this would increase the value of rejecting an offer equal to R. This would result in an increase in R and (most likely) an increase in search duration. However, a reduction in income at the 10th percentile relative to the median income would have no effect on reservation levels and search durations for women whose initial reservation levels were above the 10th percentile (this likely includes the majority of women). These asymmetric predictions can be tested in the data.

3 Data

Overview of the data: We use the India Human Development Survey 2005 (IHDS) to analyze the relationship between earnings inequality and marriage in India. The IHDS is a nationally representative survey of 41,554 rural and urban Indian households in all twenty-eight states and five union territories.[2] IHDS administered household, school, village, and medical facility surveys, collecting poverty, health, employment, economic, and social data. The breadth and depth of IHDS permits extensive analysis of human and economic development in India.

Definition of the marriage market: We define a set of spatially non-overlapping marriage markets, in which we exploit variation in the earnings distributions across markets to analyze

[2]Lakshwadeep and the Andaman and Nicobar Islands were excluded.

how earnings inequality impacts female marriage. This variation arises naturally from two facts: (1) Women tend to marry within their own caste[3], and earnings distributions tend to be very different across castes, and (2) The cultural and linguistic differences between the states of India result in marriage markets being circumscribed by state boundaries, and socio-economic differences across the states result in substantial variation in the earnings distributions. In India, about 95% of women marry within their caste (see Panel A, Figure 4). Further, after marriage, women settle very close to their place of birth: nearly 90% of married women, live within 5 hours traveling distance of their families (see Panel B, Figure 4)[4], suggesting that women marry within state, and most likely, districts.[5] The caste composition of the households in our sample is as follows: 5% Brahmins, 16% other high castes, 38% Other Backward Castes, 22% Dalits (Scheduled Castes), 6% Adivasis (Scheduled Tribes) and 12% Muslims (and a small minority of Sikh, Jain and Christians). Thus, the marriage market of a woman of "marriageable age"(as defined in Section 4) is defined as all unmarried men ages 18 to 35, who are from her same caste and state, and who are not currently enrolled in school.

Description of the earnings variables: The 2005 IHDS elicited detailed information on the employment, occupation, and earnings of individuals in the household, including the annual earnings of individuals from employment outside the home, the net income the household received from farm and business activity, and the contribution of each household individual to the farm or business, in terms of days worked in the year. From these data, we construct a comprehensive measure of annual individual earnings, which accounts for earnings from paid employment for work done outside the home and the individual's share of net income from household farm and business ventures. To obtain this share, we assign household farm and business income to each household member in proportion to the number of days an individual spent on that activity during the year (relative to the total number of days household members devoted to the activity). In our sample, about 55% of unmarried men earn income from employment solely outside the household; 34% of men receive farm income, and 13% receive some business income. In our empirical analysis we assess the robustness of the results using a narrower definition of earnings

[3]In India, caste is equivalent to a social or ethnic group. Households provide information on their caste similar to survey respondents in the US providing interviewers with race/ethnicity information.

[4]Traveling distance refers to travel by road or train, not plane.

[5]A district is an administrative division in India similar to a US county.

that only includes wage and salaried employment (from outside the home).

We use the sample weights in the IHDS to construct the earnings distribution corresponding to each marriage market. We consider only the earnings of unmarried males between ages 18 and 35, who are not currently enrolled in school. From the earnings distribution we compute (i) Mean earnings, (ii) Standard deviation of earnings, (iii) Various percentiles of the distribution, including the 90th, 50th, and 10th. In our sample, men earn close to Rs. 30,000 (approximately $680) per year, with those at the 90th percentile earning 23 times as much as men at the 10th percentile. We also derive measures of the mean and dispersion of the earnings distributions for unmarried women ages 18 to 35 who are not in school, which we utilize to test the robustness of our results.

Sample of women: Our analysis uses two overlapping samples of women. The first sample is derived from the household roster, which provides the age and marital status of each household member. We will refer to this as the "full sample". The second sample is based on the sample of ever-married women (one such woman was randomly chosen from each surveyed household), who were administered a separate questionnaire. For these women, we know their marriage histories, educational attainment, and anthropometric measurements. We refer to this sample as the "ever-married sample". Table 1 introduces this ever-married sample of 33,482 women and their households. The ever-married woman is (on average) 33 years old (standard deviation 8 years), was 17.53 years old at the time of her marriage (standard deviation 3.79), and has 2.29 children living with her in the household (standard deviation 1.44). The average number of household members is 5.2 (standard deviation 2.5), and the number of children between the ages of 0 and 14 years is 1.6 (standard deviation 1.6). The highest education level among adults (over 21 years) in the household is 7.5 years (standard deviation 5.1 years). The monthly consumption per capita is, on average, Rs. 953 (standard deviation Rs. 1,024) and the yearly income is estimated at Rs. 53,922 (standard deviation Rs. 83,284). The asset index, ranging from 0 to 30, measures household possessions and housing quality and is 12.3 (standard deviation 6.3).

4 Empirical Strategy

4.1 Baseline: Effect on female marriage

Our baseline regressions investigate the effects of male income dispersion on female marriage using a simple specification. We begin by looking at how male earnings inequality affects the probability of being married conditional on age, using the full sample of women obtained from the household roster. We employ the following linear probability model:[6]

$$y_{ics} = \alpha + \beta_1 \sigma_{ics} + \beta_2 \mu_{ics} + \gamma Age_{ics} + \eta_c + \eta_s + u_{ics} \tag{12}$$

where y_{ics} is an indicator that takes the value 1 if individual i belonging to caste c and residing in state s is married; σ_{ics}, the measure of earnings dispersion, is the standard deviation of male earnings in the marriage market defined by caste c and state s; μ_{ics} denotes the mean of the male earnings distribution in the marriage market; Age_{ics} denotes the age of the individual (thereby capturing differences in c, the utility of staying single, by age); η_c denotes a caste fixed effect; η_s denotes a state fixed effect and u_{ics} is an unobserved error term. The inclusion of caste fixed effects allows us to control for marital norms and socio-economic characteristics that differ across castes. The inclusion of state fixed effects controls for regional differences in socio-economic conditions that may impact the marriage decisions of women, as well as for regional differences in marriage norms. The standard errors are corrected by clustering at the marriage market level.

The main coefficient of interest is β_1, which represents the effect of a one unit increase in male earnings dispersion on an individual's probability of being married. The marriage market search hypothesis, as outlined in Section 2, suggests that β_1 should be negative, i.e. greater dispersion in male earnings will result in a lower per-period probability of getting married, and β_2 should be positive.

One concern with the analysis above is that the earnings dispersion may be changing rapidly over time. As a result, the proportion of married women in a given marriage market may not truly reflect the effects of the current earnings dispersion, especially if many of the married women were

[6]We opted for a linear probability model instead of a probit as the latter failed to converge due to the large number of fixed effects in several specifications. Where convergence was achieved, we obtained very similar results (results are available on request).

married several years prior to the survey. To mitigate this concern, we limit the regression sample to women between the ages of 15 and 30.

While the analysis based on specification (12) can tell us how the male earnings dispersion affects the probability of being married conditional on age, it does not directly indicate the extent to which marriage is delayed as a consequence of increased earnings dispersion. This question may be answered by turning to the sample of ever-married women for whom we know the age at marriage. For this sample, we employ the following regression specification:

$$AgeMarriage_{ics} = \alpha + \beta_1 \sigma_{ics} + \beta_2 \mu_{ics} + \eta_c + \eta_s + u_{ics} \tag{13}$$

where $AgeMarriage_{ics}$ denotes the age at which woman i belonging to caste c and residing in state s got married. In this specification, we expect β_1 to be positive, indicating that greater earnings inequality results in a later age at marriage. In line with the theoretical model presented in Section 2, we expect the coefficient β_2 to be negative.

We want to use specification (13) to examine whether women in higher-inequality markets get married later than women in lower-inequality markets. Because some of the women in the married sample were married many years ago (when earnings inequality may have been very different), it would be preferable to restrict attention to those women who were married recently. We would also prefer to exclude older women, because even though some of them may be recently married, they may have been on the marriage market for a long time and current earnings inequality may not really explain their late marriages. Our solution to these issues is to restrict the regression sample to all married women between the ages of 15 and 25 who got married within the last two years (restricting the sample further to women who were married in the last year reduces the sample drastically).[7] Note that as all women in this context get married, the sample of women who got married the last two years can be seen as a representative sample of all women, and hence there is no sample selection bias.[8]

[7]Though we recognize that social norms regarding marriage might differ across the years, Equation (13) does not include the current age of the married woman as a control. In the sample of ever-married women, age reflects search duration in the marriage market, i.e., in high inequality markets, the sample of ever-married women will be, on average, older than in low-inequality markets, or age itself is an outcome variable of the treatment "inequality". For this reason, the age of the individual is a "bad control", in the language of Angrist and Pischke (2009), because its inclusion can bias the estimate of the causal effect of male earnings inequality.

[8]To the extent that censoring may affect our results, we also estimate hazard models. Estimates from a Cox propor-

In Section 2, we noted the caveat that the typical earnings distribution is not normal, and that therefore an increase in the standard deviation while holding the mean constant does not necessarily constitute a mean-preserving spread (recall that the comparative static results in Section 2 pertained to a mean-preserving spread). For this reason, the effect of an increase in the standard deviation in the data may not necessarily correspond to the theorized effect of a mean-preserving spread. We also noted that increases in income inequality in a number of countries, including India, have largely been due to increases in the incomes of highest-earning groups. It is therefore clearly important to consider finer measures of income inequality. Following Loughran (2002), we construct measures of earnings inequality using differences between earnings at different points in the earnings distribution. Specifically, we consider the difference between earnings at the 90th and 50th percentiles, and the difference between earnings at the 50th and the 10th percentile. In Section 2 we theorized that, for most women, increases in income at the upper tail of the income distribution are likely to increase the duration of marital search, whereas reductions in income at the lower tail of the income distribution should not affect search duration. We now test these hypotheses by examining how the percentile-based differences in earnings affect female marriage, while controlling for the median earnings. The regression specifications are:

$$y_{ics} = \alpha + \beta_1(e^{90} - e^{50}) + \beta_2(e^{50} - e^{10}) + \beta_3 e^{50} + \gamma Age_{ics} + \eta_c + \eta_s + u_{ics} \tag{14}$$

$$AgeMarriage_{ics} = \alpha + \beta_1(e^{90} - e^{50}) + \beta_2(e^{50} - e^{10}) + \beta_3 e^{50} + \eta_c + \eta_s + u_{ics} \tag{15}$$

where e^{90}, e^{50}, and e^{10} denote the 90th, 50th and 10th percentiles, respectively, of the male earnings distribution in the relevant marriage market. We expect β_1 to be negative in (14) and positive in (15), while we expect β_2 to be small and not significantly different from zero in both specifications.

4.2 Extension: Effect on female education

In an extension, we hypothesize that to the extent increased male inequality delays marriage, it may also have an impact on female educational attainment by keeping women in school longer.

tional hazard model are consistent with the results from the linear specification of Equation (13). As the probability of marriage is likely to increase for older women, we also estimate a Weibull model, which allows for the hazard to monotonically increase with age. These results are also consistent with estimates from the linear model, and are available upon request.

To test this hypothesis, we employ the following specification:

$$EduYears_{ics} = \alpha + \beta_1(e^{90} - e^{50}) + \beta_2(e^{50} - e^{10}) + \beta_3 e^{50} + \eta_c + \eta_s + u_{ics} \tag{16}$$

where we use $EduYears_{ics}$, the completed educational attainment of ever-married women, as the dependent variable to determine how many extra years of education are added by greater earnings inequality. All other terms are the same as in specification 15.

Next, we attempt to determine where in the educational trajectory the additional years (if any) accrue. To do this, we construct a set of nested indicator variables for different levels of educational attainment: the first takes the value 1 if the woman received zero years of education and 0 otherwise; the second takes the value 1 if the woman did not complete primary school (i.e. received fewer than 5 years of education); the third takes the value 1 if the woman did not matriculate at middle school (i.e. received fewer than 8 years of education); the fourth takes the value 1 if the woman did not matriculate at high school (i.e. received fewer than 10 years of education); the fifth takes the value 1 if the woman did not complete high school (i.e. received fewer than 12 years of education); the sixth takes the value 1 if the woman did not complete her college education (i.e. received fewer than 15 years of education). We then estimate linear probability models on the ever-married sample, using these indicators as dependent variables:

$$level_{ics} = \alpha + \beta_1(e^{90} - e^{50}) + \beta_2(e^{50} - e^{10}) + \beta_3 e^{50} + \eta_c + \eta_s + u_{ics} \tag{17}$$

where $level_{ics}$ is an indicator for each level in the education trajectory, and e^{90}, e^{50}, and e^{10} are our measures of male income inequality.

The hypothesis of the marital search model suggests that increased search duration due to widening earnings inequality should largely impact educational attainment at the later stages of the educational trajectory because that is when women are likely to be on the marriage market, i.e. it may affect rates of high-school matriculation and completion (and even possibly college education), but is unlikely to move women from zero to positive levels of education or to raise primary school completion rates.[9] In a sense, this provides another falsification test for our hypothesis. If

[9]We do not entirely rule out the possibility, however, because there is a small percentage of women who are enrolled in school in their late teens, but have only completed four or five years of education. For these individuals an increase

greater earnings inequality is associated with a significant proportion of women completing lower levels of education, then this would suggest either that marital search duration is not the channel by which earnings inequality affects education and/or that our measures of earnings inequality are correlated with unobservable socio-economic characteristics that affect education (and possibly marriage) not accounted for in our current specifications.

We extend this logic to consider if male income inequality affects the age at marriage for women with no or little education. If in fact women are acquiring more education prior to entering the marriage market, then women with little educational attainment should not respond to changes in the male income distribution. We test this hypothesis using a variation of specification (15): we interact the male income inequality measures, $(e^{90} - e^{50})$ and $(e^{50} - e^{10})$, with an indicator for women's educational attainment. In one version the indicator is equal to 1 if the women has no education (0 years); in another version the indicator is equal to 1 if the women did not complete primary school. These variations allow us to better identify the order of events: whether women acquire more education in response to increased search duration, as the marriage search model suggests, or if educational attainment occurs before entering the marriage market. Within the our search framework, we anticipate that even among women with no or little education, their age at marriage will increase in response to greater male earnings dispersion.

Finally, we implement a falsification test that examines the schooling of girls who are not yet of marriageable age. If, for this group, greater earnings inequality is associated with more years of education and/or a greater propensity to be enrolled in school, then this would again suggest either that the channel of influence is not marital search duration or that earnings dispersion is correlated with unobserved socio-economic characteristics that affect education. To implement this test, we estimate regression specification (14), but now consider (i) Completed years of education and (ii) Current school enrollment status for unmarried girls under the age of 10 (who are not likely to be in the marriage market).

in search duration may result in primary school completion.

5 Results

5.1 Effect of income inequality on female marriage

Table 2 reports the results of the baseline regressions (12) and (13). Column 1 reports the results using the full sample (Equation (12)), while Columns 2 reports the results for the ever-married sample (Equation (13)). An increase in the standard deviation of the male earnings distribution by one unit (10,000 Rupees) reduces the probability of a woman being married by 0.5 percentage points (Column 1). However, the same change in σ_{ics} does not have a significant effect on the age at marriage (Column 2). The discrepancy between the results for the two samples may partly reflect the difference between the two samples, as well as the smaller size of the ever-married sample. More importantly, it should be recalled that unless the change in σ_{ics} represents a mean-preserving spread, the theoretical effect on search duration is ambiguous. For this reason, the results of these regressions should be interpreted with caution. An increase in the mean of the earnings distribution has a small and statistically insignificant effect in both samples. Again, this is consistent with our expectation that while an increase in the mean should technically reduce search duration (and hence increase the probability of being married), in practice the effect is likely to be very small. The coefficient on age in Column 1 indicates the probability of being married increases by 7 percentage points every year. The coefficients on the caste dummies in Columns 1 indicate that the probability of being married (conditional on age) is lowest for Brahmins (the omitted caste category), with the probability increasing as we move from other high-caste to OBC, Dalits, and Adivasis. This is consistent with the caste coefficients in Column 2: Brahmin women are, on average, three-fourths of a year older when they get married than high-caste women. The age-gap increases across the other caste categories, with the largest gap occurring between Brahmins and Adivasis (close to 1.5 years). Note that we restricted our analysis to Hindus only, and did not include Muslims, Sikh, Jain and Christians in the analysis.

We now turn to finer measures of earnings inequality. Table 3 presents the estimates from regressions (14) and (15), for the full sample (Column 1) and ever-married sample (Column 2), respectively. Because of the inherent difference in scales between the measures of upper-tail and lower-tail earnings inequality, we have converted these measures into Z-scores in order to facilitate a comparison of the coefficients on these variables.[10] The results in Table 3 paint a consistent picture. A unit increase in standardized upper-tail earnings inequality lowers the probability of getting married (at any given age) by 1.6 percentage points, and correspondingly results in a delay of marriage by about 0.34 years. However, an increase in lower-tail inequality has a small and statistically insignificant effect on both marriage propensities as well as the age at marriage.

5.2 Effect of income inequality on female education

To estimate the effect of earnings inequality on female educational attainment, we restrict ourselves to the ever-married sample, as these women have completed their education and are no longer in school. Table 4 reports the regression results for the educational attainment variables. The dependent variables in Columns 1-7 are, respectively, (i) Completed years of education, (ii) Indicator for zero years of education, (iii) Indicator for having obtained fewer than 5 years of education, (iv) Indicator for having obtained fewer than 8 years of education, (v) Indicator for having obtained fewer than 10 years of education, (vi) Indicator for having obtained fewer than 12 years of education, and (vii) Indicator for having obtained fewer than 15 years of education. A doubling of 90th percentile earnings relative to the median increases completed years of education by 0.6 years. Turning to the specific attainment variables, we find that earnings inequality has small and statistically insignificant effects on lower-level educational attainment, but it has a significant effect on high-school completion and college matriculation rates. In particular, the probability of completing high-school (Column 6) increases by 8 percentage points and the probability of entering college increases by 6 percentage points (Column 7).

One also may hypothesize that women get married later because they decided to get more education. That is, the total search duration remains the same, but women enter the marriage market later, after acquiring more education. We present results excluding this possibility in Table 5, by examining the impact of increased earnings dispersion on the age at marriage for women who completed their educational attainment prior to entering the marriage market. We find that even women who have never been to school (who have zero years of education) or who very little education (did not complete primary school) delay marriage to a greater extent as women with

[10] Define, for instance, $Z_{90-50} = \frac{(X_{90}-X_{50})-\mu_{X_{90}-X_{50}}}{\sigma_{X_{90}-X_{50}}}$.

higher educational attainment. Women with zero years of education delay marriage by nearly 0.6 years (Column 1) and women who have not completed primary school delay marriage by about 0.5 years (Column 2).

Table 6 examines the effect of earnings inequality on educational attainment and current school enrollment of unmarried children below the age of 10. Because these children are plausibly not on the marriage market, earnings inequality should not affect these educational outcomes. The results in Table 6 confirm this hypothesis: neither current school enrollment nor completed years of education are correlated with male earnings inequality for this sample.

These results are consistent with the hypothesis that greater earnings inequality affects female marriage by increasing search duration, and therefore impacts educational attainment at the point of the educational trajectory that women find themselves at when they enter the marriage market. In the next section, we test these results against a number of alternative hypotheses.

6 Testing alternative hypotheses

We have established that in high-inequality markets, women tend to postpone marriage. By including state and caste fixed effects in our regressions in Table 3, we attempt to control for unobservable differences between women in high- and low-inequality markets that may arise from socio-economic conditions specific to regions or societal groups. Nonetheless, the main threat to identification of the causal effects remains that earnings inequality may be correlated with unobserved confounding factors that may also affect marriage. In this section, we address this potential omitted variable bias.[11]

We first show that the effects found in Table 3 are not symmetric, i.e., men's probability of being married does not relate to women's earnings inequality (Table 7). We then demonstrate that women in marriage markets with high inequality are on average not different from women in low inequality markets, in terms of characteristics such as height and the age at which they reached menarche (Table 8). As neither of these variables represents a short-term decision that should be affected by earnings inequality, they are suitable for testing for differences in observable

[11] We also subject the results in Table 3 to a range of robustness tests, using alternate earnings measures and alternate regression samples. The impact of male income inequality on female marital outcomes remains the same. These results are available upon request.

characteristics between women in high- and low-inequality markets. Finally, we provide evidence that the results are not due to (i) Increased duration of labor market search on the part of men as a response to greater dispersion, (ii) Male earnings dispersion proxying for expected or current female earnings, or (iii) Financial constraints faced women's families (Tables 9A and 9B).

It is possible that the results presented in Table 3 could reflect increased male search duration due to increased female earnings inequality, and if male earnings dispersion is correlated with female earnings dispersion, then this would appear as a negative correlation between the female marriage age and male income inequality. To investigate this possibility, we re-estimate Equation (14), now considering the effect of female income inequality on the probability of male marriage, in a sense a falsification test of our original specification. Table 7 presents the results. We find that is no symmetric effect: the coefficients for the measures of the upper-and lower-tail of the female earnings distribution are both small and statistically insignificant. Thus men are not searching longer due to increased dispersion in female earnings.

We note that woman in high-and low-inequality markets could be different from each other, resulting in their different marriage patterns. In this section, we verify whether women in high- and low-inequality markets are similar in observable characteristics and whether inclusion of these characteristics changes our results. Specifically, we look at the correlation between the earnings inequality measures and two outcomes that are not decision variables, namely the height of the woman and the age at which she attained menarche, which are reported for the ever-married women sample. Because neither of these variables represents a short-term decision that should be affected by earnings inequality, they are suitable for testing for differences in observable characteristics between women in high- and low-inequality markets. Both height and age at menarche reflect childhood nutrition, however, so they may be correlated with earnings inequality if the latter is correlated with unobserved socio-economic characteristics. The variable age at menarche is of particular interest because women typically get married only after they attain menarche (indeed, in our sample of ever-married women, age at menarche is a very strong predictor of age at marriage). If individuals in marriage markets with greater earnings inequality also happen to be associated with poor childhood nutrition (which would act to delay menarche), this might result in a spurious positive correlation between earnings inequality and age at marriage. The height of the woman, in addition to being a correlate of socio-economic status, may also directly impact the age at marriage (e.g. because men typically marry women shorter than themselves, it is plausible that tall women may have to search longer for a suitable groom).

Table 8 presents the results of examining whether male earnings inequality predicts age at menarche (Column 1) and height of women (Column 2) in the ever-married sample. The results are negative: the coefficients on the earnings inequality variables are small and statistically insignificant, confirming that women in high inequality markets are not different from women in low inequality markets in terms of observable characteristics. In accordance with these results, including age at menarche and height as controls in regression (15) does not significantly affect the estimated effect of male earnings inequality on age at marriage, as shown in Column 3 of Table 8.

Table 9A reports the regression results of equation (14) using the full sample, with the following additional controls included: male/female sex-ratio (defined as the number of unmarried (eligible) men to the number of unmarried (eligible) women) (Column 2), female earnings dispersion (Column 3), and mean expense the bride's family usually incurs for a wedding (Column 4). Table 9B reports the corresponding regression results for the ever-married sample (estimating Equation 15).

The inclusion of sex-ratio is motived as follows. It is possible that increasing the dispersion of the male earnings distribution may extend the labor market search of men and thereby delay their own entry into the marriage market (or make them ineligible for marriage). This may reduce the number of available men in the marriage market, possibly delaying female marriage. Alternatively, it is plausible that men only join the marriage market when they reach a minimum earnings level. An increased dispersion of earnings, could increase the number of men on the market in this case, affecting the probability of marriage for women. Finally, it might be possible that men in high-inequality areas move to areas with low-inequality, hence reducing the number of available men in high-inequality areas. Column 2 in Tables 9A and 9B show that inclusion of sex-ratio has little impact on the coefficients of the earnings inequality variables. We note however that the sex-ratio is also likely to be endogenous and its coefficient should not be interpreted causally.

It may also be the case that inequality in the upper-tail of the male earnings distribution captures the skill/education premia, and these returns to higher education may motivate women or men to pursue higher education (delaying their marriage). To address this we control for measures of female earnings dispersion (Column 3 of Tables 9A and 9B). We note, however, that the female earnings variables are potentially endogenous. Nonetheless, our intention is not to interpret the coefficients on these variables, but to examine whether their inclusion affects the coefficients of the male earnings inequality variables. The results are qualitatively the same as the baseline results of Table 3, although the estimated effect of earnings inequality on the age at marriage is now generally larger than in previous specifications.

We also consider the possibility that increased male earnings dispersion at the upper-tail may actually induce matches to form more quickly if women and their families are making offers to secure high-earning men. However, the realization of the match, i.e. the marriage, may be delayed as it takes time to afford the wedding, especially if more attractive candidates ask for greater dowries, thereby resulting in a spurious positive correlation between earnings inequality and age at marriage. To address this, we control for the average expenditure the bride's family incurs for a wedding. The inclusion of wedding expenses has little impact on the coefficients on the earnings inequality measures (Column 4 of Tables 9A and 9B). Again we note the caveat that marriage expenditures may be endogenous and its coefficient should not be interpreted causally.

7 Conclusion

We use a nationally representative dataset from India to test whether women delay their marriage as a response to increased earnings inequality of men. In line with the predictions from a marital search model, we find that increases in upper tail inequality delay marriage, while increases in lower tail inequality have no significant effect. There is a corresponding effect on educational attainments, with women in high-inequality markets being more likely to obtain higher education. Marriage is delayed even for women who attained no education, ruling out the hypothesis that greater inequality may delay marriage by causing women to seek higher education, i.e. the direction of effect runs from marriage to education, not the other way around.

In particular, we find that a unit increase in standardized upper-tail earnings inequality lowers the propensity to get married (at any given age) by 1.6 percentage points, and correspondingly results in a delay of marriage by about 0.34 years. Comparing these results with the US context: Gould and Paserman (2003) using US census data from 1970 till 1990 find that higher male inequality in a city lowers the marriage rate of women, and that this effect is not unimportant:

increasing male inequality explains about 30 percent of the marriage rate decline for women of the last few decades. Loughran (2002), using the same data, finds that rising within-group make wage inequality accounts for anywhere between 5 and 35 percent of the decline in the age-specific female propensity to marry between 1970 and 1990, the strongest results being for women age 22-28 with less than a college education.

The richness of the data allows us to test and rule out a number of alternative hypotheses for these findings. The results are not due to (i) Men searching longer in the marriage market in response to greater female earnings inequality, (ii) Regional or caste-base social norms, (iii) Men searching longer in the labor market (reducing the gender ratio in the marriage market), (iv) Earnings dispersion proxying for educational premia, and hence encouraging women to stay in school longer, or (v) Women's families needing more time to afford greater dowries. While we have little information on the woman's natal family in the data, e.g. we cannot control for variables such as natal family composition and wealth, we do establish that women in high and low-inequality markets are not different with respect to some observable characteristics: height and age at menarche.

There are clear long-term implications of the phenomenon we have documented. The acquisition of more education due to increased search duration on the marriage market might increase the future income stream of women, contributing to their economic well-being. In addition, greater educational attainment may improve their bargaining power in the household, allowing women to direct resources towards children's, especially girls', education and health, thus affecting the skills and productivity of the future workforce.

In ongoing research, we are using data from the National Sample Surveys (NSS) to determine the share of the increase in age at marriage over the last two decades that can be attributed to increasing inequality. The NSS data would address some of these cohort concerns as one could match women who were young at a particular time with wages at that time.

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Appendix (Proofs)

We provide outlines for the proofs of the results cited in Section 2.2. We showed that the reservation earnings R could be obtained by solving the equation:

$$[r+1-F(R)]R = rc + \int_R^{\infty} x dF(x) \tag{18}$$

where $r = \frac{1-\beta}{\beta}$.

We now assume that the F distribution is normal with mean μ and standard deviation σ. We can now apply the standard formula for the mean of a truncated normal distribution to write:

$$\begin{aligned} \int_R^{\infty} x dF(x) &= P(x > R) \cdot E(x|x > R) = P(x > R) \cdot \left[\mu + \sigma \frac{\phi\left(\frac{R-\mu}{\sigma}\right)}{1 - \Phi\left(\frac{R-\mu}{\sigma}\right)} \right] \\ &= \mu \left[1 - \Phi\left(\frac{R-\mu}{\sigma}\right) \right] + \sigma\phi\left(\frac{R-\mu}{\sigma}\right) \end{aligned} \tag{19}$$

Thus, we have, from (18):

$$\left[r + 1 - \Phi\left(\frac{R-\mu}{\sigma}\right) \right] R = rc + \mu \left[1 - \Phi\left(\frac{R-\mu}{\sigma}\right) \right] + \sigma\phi\left(\frac{R-\mu}{\sigma}\right) \tag{20}$$

We now assume that the initial (starting) distribution of F is $N(0,1)$. This simplifies some of the algebra. Using Equation (20) above, and noting that $\phi'(z) = -z\phi'(z)$, one can obtain the following derivatives:

$$\frac{dR}{dc} = \frac{r}{q} > 0 \tag{21}$$

$$\frac{dR}{d\mu} = \frac{q}{r+q} < 1 \tag{22}$$

$$\frac{dR}{d\sigma} = \frac{\phi(R)}{r+q} > 0 \tag{23}$$

These were the comparative static results presented in Section 2.2.

Figure 1: Average Age at Marriage (in India), 1980-2005

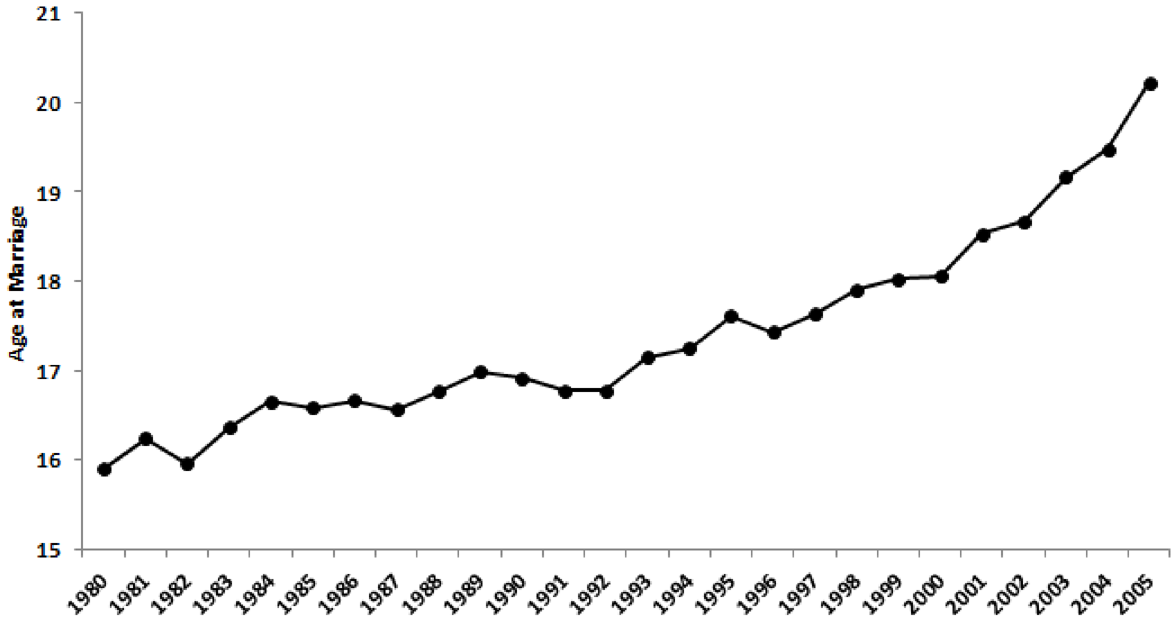

Figure 1 depicts the average age at marriage for women in India, from 1980-2005. Over this period, we see an increase in the marriage age from 16.25 years to 20.36 years.
Source: Authors' calculation, IHDS data.

Figure 2: Completed Years of Education, 1980-2005

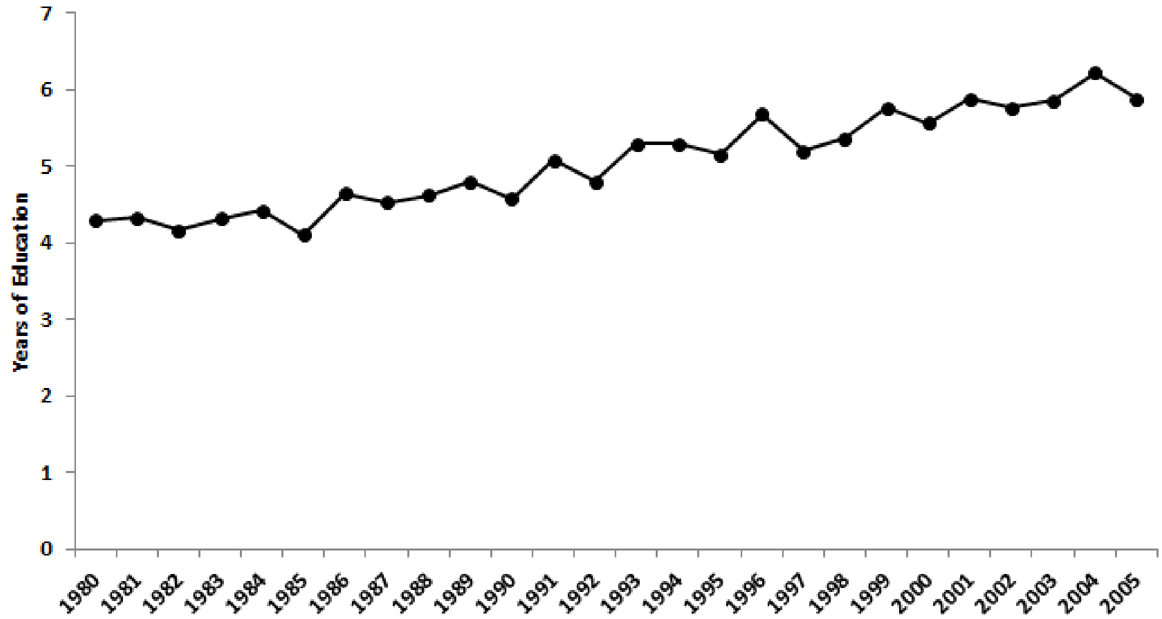

Figure 2 depicts the average years of educational attainment amongst married women in India, from 1980-2005. Over this period, we see an increase in the years of education from 4.65 to 6.47. We note that this increase in educational attainment does not fully explain the four year increase in the marriage age of women depicted in Figure 1.
Source: Authors' calculation, IHDS data.

Figure 3: Functional form of $R - \frac{dR}{d\sigma}$ as a function of R (reservation level) (in red)

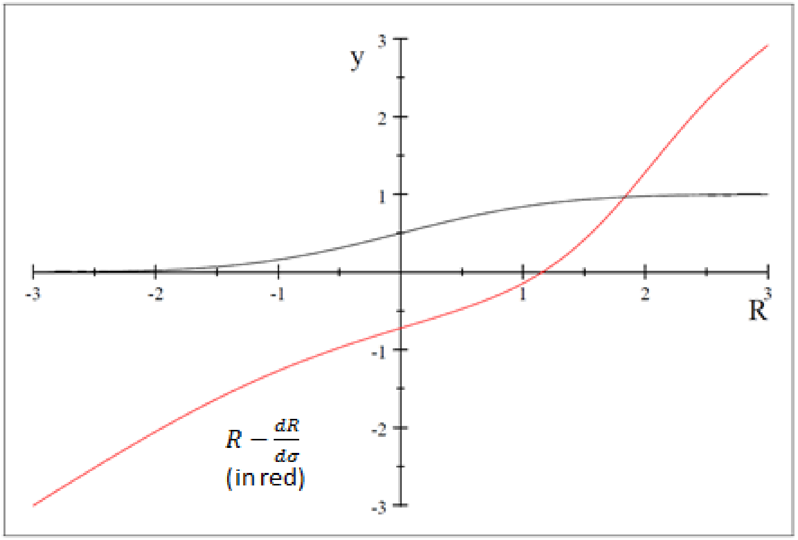

Figure 3 graphs $R - \frac{dR}{d\sigma}$ against different values of R (where we have assumed a discount rate of $\beta = 0.95$). The figure also graphs the standard normal cumulative distribution function. The figure shows that $R - \frac{dR}{d\sigma}$ is negative over a substantial range of values of R, implying that increasing earnings dispersion most likely increases search duration and the age at marriage.

Figure 4: Marriage Market Characteristics

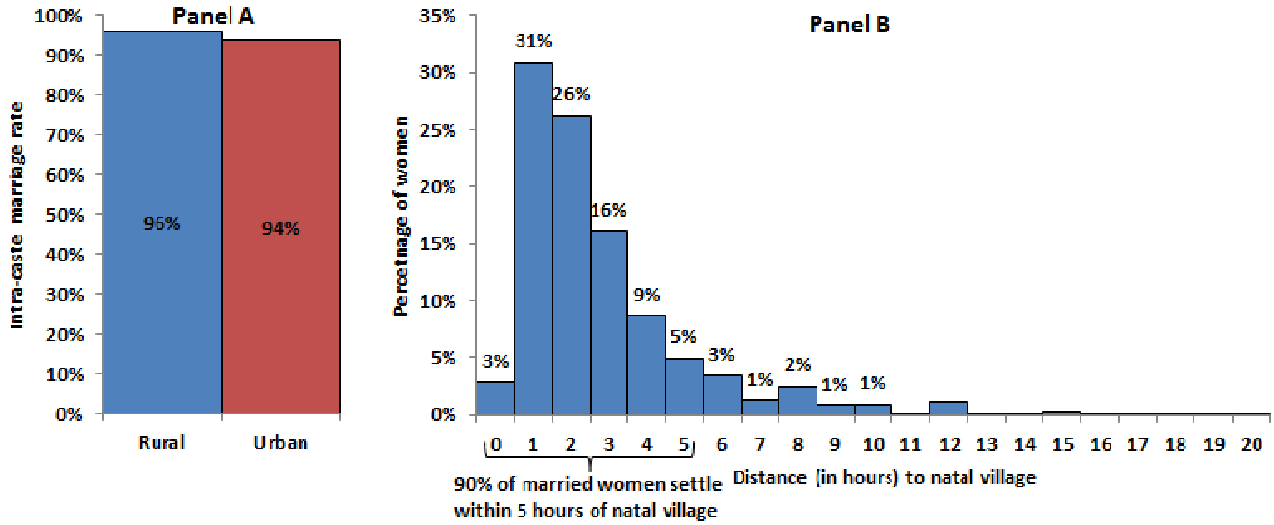

Figure 4 depicts the marriage market in India. Panel A shows that the intra-caste marriage rate is high: 96% in rural areas, and 94% in urban areas. Panel B shows that after marriage, women in India settle very close to their place of birth. (Traveling distance refers to travel by road or train, not plane.) Combined, this provides justification to define the marriage market according to a woman's caste and geographical location.
Source: Authors' calculation, IHDS data.

Table 1: Introducing the "ever-married" women sample

	Mean	Std. Dev.
Age of ever-marrried woman (years)	33.0	8.0
Age at marriage of ever-married woman (years)	17.53	3.79
Number of children of ever-married woman	2.29	1.44
Number of household members	5.2	2.5
Number of children (0-14 years) in the household	1.6	1.6
Years of education (adult $>$ 21 years)	7.5	5.1
Monthly Consumption per capita (in Rs)	953	1,024
Income (yearly, in Rs)	53,922	83,284
Asset index (range 0 to 30)	12.3	6.3

Table 1 presents descriptive statistics on the ever-married sample of women. The household is defined as the persons who live under the same roof and share the same kitchen for 6+ months. Asset index: IHDS asked a series of questions about what goods the household owned and about the quality of the housing. This index sums 30 dichotomous items measuring household possessions and housing quality. The age of marriage is noted for women who are married just once. The number of children only includes the children of the woman living in the household.
Source: Authors' calculation, IHDS data.

Table 2: Effect of Male Earnings Inequality (standard deviation and mean) on Female Marriage

	(1) Probability of marriage	(2) Age at marriage
Standard deviation of male earnings	-0.005***	-0.012
	(0.001)	(0.034)
Mean of male earnings	-0.005	0.231
	(0.008)	(0.227)
Age of woman (years)	0.068***	
	(0.001)	
(Other) High-caste indicator	0.048***	-0.761*
	(0.013)	(0.426)
OBC indicator	0.078***	-1.112**
	(0.014)	(0.496)
Dalit indicator	0.104***	-1.303**
	(0.016)	(0.515)
Adivasi indicator	0.114***	-1.509**
	(0.021)	(0.742)
Constant	-0.946***	20.031***
	(0.032)	(0.854)
State fixed effects included?	Yes	Yes
N (women)	25,549	645
R-squared	0.451	0.165

Table 2 reports the results of the baseline regressions. Column 1 reports the results using the full sample (dependent variable: indicator of marital status, Linear Probability Model), while Column 2 reports the results for the ever-married sample (dependent variable: age of marriage in years). We find that an increase in the standard deviation of the male earnings distribution by one unit (10,000 Rupees) reduces the probability of a woman being married by 0.5 percentage points (Column 1), but does not have a significant effect on the age at marriage (Column 2). Notes: The excluded caste category is Brahmins; Standard errors are clustered by marriage market; * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table 3: Effect of Male Earnings Inequality (upper and lower tails) on Female Marriage

	(1) Probability of marriage	(2) Age at marriage
Difference in male earnings: 90^{th}-50^{th} percentile	-0.016***	0.343***
	(0.006)	(0.109)
Difference in male earnings: 50^{th}-10^{th} percentile	-0.007	0.055
	(0.009)	(0.287)
Male earnings at 50^{th} percentile	0.014	-0.476**
	(0.013)	(0.233)
Age of woman (years)	0.068***	
	(0.001)	
(Other) High-caste indicator	0.043***	-1.056***
	(0.014)	(0.394)
OBC indicator	0.108***	-2.037***
	(0.016)	(0.502)
Dalit indicator	0.138***	-2.258***
	(0.016)	(0.461)
Adivasi indicator	0.156***	-2.651***
	(0.021)	(0.686)
Constant	-1.017***	21.967***
	(0.031)	(0.706)
State fixed effects included?	Yes	Yes
N (women)	25,550	646
R-squared	0.451	0.174

Table 3 reports the effect of the difference between earnings at the 90th and 50th percentile and the difference between earnings at the 50th and the 10th percentile on the marriage of women. Column 1 reports the results using the full sample (dependent variable: indicator of marital status, Linear Probability Model), while Column 2 reports the results for the ever-married sample (dependent variable: age of marriage in years). Because of the inherent difference in scales between the measures of upper-tail and lower-tail earnings inequality, we converted these measures into Z-scores (Z score $= \frac{x-\mu}{\sigma}$) in order to facilitate a comparison of the coefficients on these variables. The results in Table 3 paint a consistent picture. A unit increase in standardized upper-tail earnings inequality lowers the propensity to get married (at any given age) by 1.6 percentage points, and correspondingly results in a delay of marriage by about 0.34 years. However, an increase in lower-tail inequality has a small and statistically insignificant effect on both marriage propensity and the age at marriage. Notes: The excluded caste category is Brahmins; Standard errors are clustered by marriage market; * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table 4: The effect of male earnings inequality on female educational attainment (ever-married sample)

	(1) Years of education	(2) No education	(3) Under 5 years	(4) Under 8 years	(5) Under 10 years	(6) Under 12 years	(7) Under 15 years
Difference in male earnings: 90^{th}-50^{th} percentile	0.607*	-0.013	-0.016	-0.052	-0.042	-0.076***	-0.057***
	(0.313)	(0.026)	(0.027)	(0.035)	(0.029)	(0.027)	(0.021)
Difference in male earnings: 50^{th}-10^{th} percentile	0.270	-0.052	-0.035	0.042	-0.009	-0.057	0.010
	(0.590)	(0.047)	(0.049)	(0.066)	(0.060)	(0.058)	(0.053)
Male earnings at 50^{th} percentile	-0.656	0.124**	0.104**	0.068	0.065	0.041	-0.099*
	(0.541)	(0.051)	(0.048)	(0.047)	(0.045)	(0.057)	(0.056)
High-caste indicator	-1.901**	0.138**	0.118**	0.170**	0.206***	0.063	0.080
	(0.776)	(0.057)	(0.057)	(0.066)	(0.072)	(0.091)	(0.096)
OBC indicator	-4.446***	0.319***	0.293***	0.328***	0.466***	0.296***	0.179
	(0.931)	(0.075)	(0.076)	(0.087)	(0.086)	(0.112)	(0.109)
Dalit indicator	-5.535***	0.366***	0.374***	0.460***	0.581***	0.363***	0.191*
	(0.882)	(0.077)	(0.077)	(0.082)	(0.079)	(0.108)	(0.105)
Adivasi indicator	-6.507***	0.577***	0.524***	0.482***	0.538***	0.287*	0.164
	(1.322)	(0.107)	(0.105)	(0.111)	(0.113)	(0.153)	(0.132)
Constant	12.097***	-0.242*	-0.161	0.002	0.148	0.515***	0.918***
	(1.374)	(0.128)	(0.120)	(0.120)	(0.115)	(0.161)	(0.160)
State fixed effects included?	Yes	Yes	Yes	Yes	Yes	Yes	Yes
N (women)	627	627	627	627	627	627	627
R-squared	0.201	0.139	0.130	0.164	0.176	0.167	0.232

Table 4 reports the regression results for the educational attainment variables. The dependent variables in Columns 1-7 are, respectively, (i) Completed years of education, (ii) Indicator for zero years of education, (iii) Indicator for having obtained fewer than 5 years of education, (iv) Indicator for having obtained fewer than 8 years of education, (v) Indicator for having obtained fewer than 10 years of education, (vi) Indicator for having obtained fewer than 12 years of education, and (vii) Indicator for having obtained fewer than 15 years of education. Columns 2-7 estimate a Linear Probability Model. We find that a doubling of 90th percentile earnings relative to the median increases completed years of education by 0.6 years and that the probability of completing high-school increases by 8 percentage points and the probability of entering college increases by 6 percentage points. Notes: The excluded caste category is Brahmins; Standard errors are clustered by marriage market; * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. See also notes to Table 3.

Table 5: Do women with no or little education experience delayed marriage?

	Age at marriage	
	(1)	(2)
Difference in male earnings: 90^{th}-50^{th} percentile	0.311***	0.311**
	(0.117)	(0.119)
Difference in male earnings: 50^{th}-10^{th} percentile	-0.017	-0.076
	(0.283)	(0.291)
Male earnings at 50^{th} percentile	-0.377	-0.348
	(0.231)	(0.236)
Female no education indicator	-0.868***	
	(0.224)	
Female less than primary indicator		-0.934***
		(0.205)
Male 90^{th}-50^{th} * No education	0.282	
	(0.239)	
Male 50^{th}-10^{th} * No education	0.097	
	(0.179)	
Male 90^{th}-50^{th} * Less than primary		0.183
		(0.217)
Male 50^{th}-10^{th} * Less than primary		0.287*
		(0.169)
Constant	21.854***	21.839***
	(0.710)	(0.720)
Net effect: 90^{th}-50^{th}	0.593**	0.494**
	(0.236)	(0.209)
State fixed effects included?	Yes	Yes
N (women)	627	627
R-squared	0.202	0.208

Table 5 examines whether women who have no or little education still delay marriage in response to greater male earnings dispersion. We report the results for a variation of the baseline age at marriage regression: the male income inequality measures are interacted with indicators for low levels of female educational attainment (Column 1: no education, Column 2: did not complete primary). Women with zero years of education delay marriage by nearly 0.6 years (Column 1) and women who have not completed primary school delay marriage by about 0.5 years (Column 2). This rules out the hypothesis that greater inequality may delay marriage by causing women to seek higher education, i.e. the direction of effect runs from marriage to education, not the other way around. Notes: The excluded caste category is Brahmins; Standard errors are clustered by marriage market; * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. See also notes to Table 3.

Table 6: The effect of male earnings inequality on educational attainment (girls under 10 years)

	(1) Currently enrolled in school	(2) Completed years of education
Difference in male earnings: 90^{th}-50^{th} percentile	0.011	0.023
	(0.012)	(0.017)
Difference in male earnings: 50^{th}-10^{th} percentile	0.023	0.001
	(0.018)	(0.039)
Male earnings at 50^{th} percentile	-0.045*	-0.022
	(0.024)	(0.054)
Age of woman (years)	0.219***	0.389***
	(0.003)	(0.014)
High-caste indicator	-0.089***	-0.123**
	(0.032)	(0.056)
OBC indicator	-0.199***	-0.259***
	(0.038)	(0.076)
Dalit indicator	-0.275***	-0.321***
	(0.040)	(0.076)
Adivasi indicator	-0.371***	-0.513***
	(0.049)	(0.094)
Constant	-0.892***	-0.711***
	(0.065)	(0.153)
State fixed effects included?	Yes	Yes
N (girls)	19,446	19,446
R-squared	0.552	0.571

Table 6 examines the effect of earnings inequality on educational attainment and current school enrollment of unmarried children below the age of 10. Column 1 has dependent variable currently enrolled (and estimates a Linear Probability Model) and Column 2 has dependent variable completed years of education. Table 6 shows that neither current school enrollment nor completed years of education are correlated with male earnings inequality for this sample. Notes: The excluded caste category is Brahmins; Standard errors are clustered by marriage market; * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. See also notes to Table 3.

Table 7: The effect of female earnings inequality on male marriage

	(1) Probability of marriage
Difference in female earnings: 90^{th}-50^{th} percentile	-0.001
	(0.004)
Difference in female earnings: 50^{th}-10^{th} percentile	-0.004
	(0.008)
Female earnings at 50^{th} percentile	0.020^{*}
	(0.012)
Age of man (years)	0.047^{***}
	(0.000)
High-caste indicator	0.030^{**}
	(0.012)
OBC indicator	0.080^{***}
	(0.010)
Dalit indicator	0.112^{***}
	(0.011)
Adivasi indicator	0.123^{***}
	(0.011)
Constant	-0.798^{***}
	(0.017)
State fixed effects included?	Yes
N (men)	37,841
R-squared	0.546

Table 7 examines the effect of female earnings inequality on the probability of men to be married using a Linear Probability Model. Table 7 shows that there is no symmetric effect: the coefficients for the measures of the upper- and lower-tail of the female earnings distribution are both small and statistically insignificant. Notes: The excluded caste category is Brahmins; Standard errors are clustered by marriage market; * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. See also notes to Table 3.

Table 8: Do pre-marriage market characteristics correlate with male earnings dispersion?

	(1) Age at menarche	(2) Height	(3) Age at marriage
Difference in male earnings: 90^{th}-50^{th} percentile	0.035	1.106	0.335^{***}
	(0.080)	(1.310)	(0.106)
Difference in male earnings: 50^{th}-10^{th} percentile	0.153	-0.831	0.054
	(0.179)	(3.878)	(0.289)
Male earnings at 50^{th} percentile	0.091	1.590	-0.489^{**}
	(0.237)	(3.706)	(0.240)
High-caste indicator	-0.061	-4.021	-1.030^{**}
	(0.397)	(4.832)	(0.404)
OBC indicator	-0.114	0.753	-2.037^{***}
	(0.563)	(6.659)	(0.518)
Dalit indicator	-0.233	2.116	-2.261^{***}
	(0.533)	(6.185)	(0.477)
Adivasi indicator	-0.012	9.290	-2.706^{***}
	(0.642)	(8.058)	(0.698)
Age at menarche			0.041
			(0.060)
Height			0.006^{**}
			(0.002)
Constant	13.961^{***}	142.825^{***}	20.540^{***}
	(0.813)	(8.961)	(1.149)
State fixed effects included?	Yes	Yes	Yes
N (women)	646	646	646
R-squared	0.235	0.086	0.180

Table 8 examines whether male earnings inequality predicts age at menarche and height of women (Columns 1 and 2) and whether including these variables into the analysis of age at marriage alters the results (Column 3). Table 8 shows that the coefficients on the earnings inequality variables are small and statistically insignificant, confirming that women in high inequality markets are not different from women in low inequality markets in terms of observable characteristics, and that including age at menarche and height as controls in regression does not significantly affect the estimated effect of male earnings inequality on age at marriage. Notes: The excluded caste category is Brahmins; Standard errors are clustered by marriage market; * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. See also notes to Table 3.

Table 9: Testing alternative hypotheses

	Probability of marriage			
	(1)	(2)	(3)	(4)
Male earnings: 90^{th}-50^{th}	-0.016***	-0.014**	-0.021***	-0.013**
	(0.006)	(0.005)	(0.007)	(0.006)
Male earnings: 50^{th}-10^{th}	-0.007	0.002	-0.012	-0.005
	(0.009)	(0.008)	(0.009)	(0.009)
Male earnings: 50^{th}	0.014	0.002	0.010	0.011
	(0.013)	(0.011)	(0.013)	(0.013)
Age of woman (years)	0.068***	0.068***	0.068***	0.068***
	(0.001)	(0.001)	(0.001)	(0.001)
Male:Female ratio		0.044***		
		(0.011)		
Female earnings: 90^{th}-50^{th}			0.012**	
			(0.006)	
Female earnings: 50^{th}-10^{th}			-0.006	
			(0.010)	
Female earnings: 50^{th}			0.017	
			(0.014)	
Wedding expenditure				-0.003
				(0.002)
Constant	-1.017***	-1.067***	-1.018***	-0.962***
	(0.031)	(0.033)	(0.032)	(0.047)
Caste fixed effects included?	Yes	Yes	Yes	Yes
State fixed effects included?	Yes	Yes	Yes	Yes
N (women)	25550.000	25530.000	25550.000	25550.000
R-squared	0.451	0.451	0.451	0.451

(a) Full sample

	Age at marriage			
	(1)	(2)	(3)	(4)
Male earnings: 90^{th}-50^{th}	0.343***	0.339***	0.467***	0.386***
	(0.109)	(0.109)	(0.139)	(0.129)
Male earnings: 50^{th}-10^{th}	0.055	0.048	0.191	0.066
	(0.287)	(0.296)	(0.312)	(0.281)
Male earnings: 50^{th}	-0.476**	-0.460*	-0.328	-0.506**
	(0.233)	(0.245)	(0.274)	(0.230)
Male:Female ratio		-0.239		
		(0.332)		
Female earnings: 90^{th}-50^{th}			-0.283	
			(0.182)	
Female earnings: 50^{th}-10^{th}			0.199	
			(0.321)	
Female earnings: 50^{th}			-0.581	
			(0.527)	
Wedding expenditure				-0.041
				(0.076)
Constant	21.967***	22.326***	21.967***	22.665***
	(0.706)	(0.897)	(0.861)	(1.417)
Caste fixed effects included?	Yes	Yes	Yes	Yes
State fixed effects included?	Yes	Yes	Yes	Yes
N (women)	646.000	644.000	646.000	646.000
R-squared	0.174	0.174	0.179	0.175

(b) Ever-married sample

Table 9 reports the regression results using the full sample (9a) (dependent variable: indicator of marital status, Linear Probability Model) and the ever-married sample (9b) (dependent variable: age of marriage), with the following additional controls included: male/female sex-ratio (defined as the number of unmarried (eligible) men to the number of unmarried (eligible) women) (Column 2), female earnings dispersion (Column 3), and the average amount the bride's family spends for a wedding (by marriage market) (Column 4). The results are qualitatively the same as the baseline results of Table 3. Notes: Standard errors are clustered by marriage market; * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. See also notes to Table 3.