Age at Marriage and Women's Labor Market Outcomes in India[*]

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Abstract

We examine the relationship between women's age at marriage and their labor market outcomes using nationally representative household data from India. Employing an instrumental variables-based empirical strategy, we find that a delay in women's age at marriage has no significant causal effect on their labor market outcomes. This is despite marriage delay being associated with higher education, lower fertility, and (possibly) higher dowry for Indian women. We argue that this might be because, older brides, as compared to younger brides, face more backlash from their partners. This backlash effect could be offsetting the positive labor market effects of marriage delay.

JEL: J12, J16, J22, J31, O12
Keywords: Age at Marriage, India, Labor Market Outcomes, Male Backlash, Women.

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Abstract

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1 Introduction

Early marriage is one of the most pressing issues currently in India. According to a report of the Advocacy group ActionAid around 33% child marriages in the world occur in India.[1] As per the report, nearly 85.2 million Indian women living as on March 1, 2011 were married before reaching the age of 18. Equally alarming is women's dismal labor market prospects in India. According to the India's National Sample Survey (NSS), only 33% of adult Indian women had a job or were actively looking for one in 2011. According to an article published in *The Economist* (July 5, 2018), as of 2018, only Saudi Arabia has a labor force participation rate of women that is lower than that of India among the G20 countries.[2] The earnings and wages of Indian women who are working are also notoriously low. According to the Global Wage Report 2018-19 published by the International Labour Organization (ILO), hourly wages of women are 34% less than men in India. This gap in wages is highest among 73 countries studied in the report.

In this paper, we examine whether these two economic problems are inter-related. Specifically we ask: can early marriage be held responsible for dismal labor market prospects of Indian women? In theory, early marriage can hamper labor market prospects of women in multiple ways. First, early marriage can negatively affect labor market prospects by interrupting the accumulation of their formal education and labor market skills (Field and Ambrus, 2008). Second, early marriage implies early motherhood which increases the value of women's time at home. This might cause younger brides to focus more on home production and reduces their likelihood of participation and productivity in the labor market (Wang and Wang, 2017). Third, early marriage might serve as a conduit for transmission of traditional norms discouraging female employment (Asadullah and Wahhaj, 2019). This might adversely affect labor market outcomes of younger brides. Finally, in countries where the system of dowry is prevalent, early marriage might be associated with lower dowry because

[1] https://www.actionaidindia.org/wp-content/uploads/2018/06/Eliminating-Child-Marriage-in-India.pdf
[2] https://www.economist.com/leaders/2018/07/05/why-india-needs-women-to-work

the demand of youthful brides is generally substantially higher than the demand for older brides (Field and Ambrus, 2008). If younger brides bring less assets to the spousal household as dowry, this might negatively influence their intrahousehold decision power (Brown, 2009) and, in turn, hamper their labor market prospects.

We use data from the Indian Human Development Survey 2012 (Desai et al., 2015), a nationally representative household survey, and focus on four types of labor market outcomes of women, namely, labor market participation, hourly earnings, annual wage earnings, and work days per year. The main empirical challenge in identifying the causal effect of age at marriage on women's labor market outcomes is that marriage age may be endogenous due to self-selection into marriages. As per Bergstrom and Bagnoli's (1993) theory based on asymmetric information, information on the earnings capabilities of individuals is available only at later stages of the life cycle. Consequently, parents of girls with high earnings potential or those parents who are more focused about their daughter's career might be more likely to be interested in delaying marriages (perhaps due to potential interruption of accumulation of both labor market skills and formal education from early marriages) until their daughter's earnings potential is fully revealed.

To address the potential issue of endogeneity, we employ an empirical strategy proposed by Field and Ambrus (2008), who instrument women's age at marriage by their age at menarche. This instrument is motivated by the observation that has been made by sociologists and anthropologists that parents become extremely anxious to get their daughters married once they have reached menarche, partly to avert any unwanted pregnancies (Caldwell et al. 1983; Srinivas 1984). Since a significant portion of the variation in timing of menarche is random (Field and Ambrus, 2008), it is likely to serve as a good instrument for the age at marriage. This instrument has been recently used by Sekhri and Debnath (2014), Chari et al. (2017), Asadullah and Wahhaj (2019), and Sunder (2019) among a few others.

Our findings are as follows. The ordinary least squares (OLS) results indicate that a year of delayed marriage of women is associated with significant improvements in their labor market outcomes. However, when we use the instrumental variable (IV) approach to examine whether these results are causal or not, we find that the age at marriage of women has no significant impact on their labor market outcomes. Further, we find that the magnitude of the IV estimates are substantially smaller than the OLS estimates. Our results are extremely robust: they do not change when we restrict our analytic sample to include only the potential compliers, when we use alternative definitions of women's labor market participation, when we carry out our analysis separately for culturally different regions, as well as when we address selection of women into labor force.

Our results are quite surprising because previous studies find compelling evidence that, for Indian women, a delay in age at marriage leads to an increase in their education and dowry payments at marriage, and a reduction in their fertility (Chowdhury, 2010; Sekhri and Debnath, 2014; Chari et al., 2017). In fact, for our sample as well, we find that a delay in women's age at marriage is clearly associated with higher education and lower fertility. As such older brides, compared to younger brides, should have had better labor market outcomes in India. Why does not that happen?

We offer a possible explanation based on the sociological theory of male backlash. According to the male backlash theory, education or empowerment might make women more vulnerable to violence from their spouses and in-laws in a patriarchal society (Macmillan and Gartner, 1999). This is because women's education or empowerment threatens male dominance and female dependence – the fundamental feature of patriarchy. Since in our sample, older brides are indeed more educated (and hence more empowered) than the younger brides, it might be the case that the older brides face more violence from their spouses and in-laws compared to younger brides. This could be negatively affecting the labor market participation and productivity of the older brides, which, in turn, could be offsetting the positive effects of marriage delay on their labor market prospects through better education, higher dowry and lower fertility. We discuss this explanation in detail towards the end of this paper.

1.1 Related Literature

Our study is related to the literature on socioeconomic effects of women's age at marriage. Previous studies have examined the effects of women's age at marriage on various outcomes including own education, health, attitudes towards traditional gender norms and labor market outcomes, as well as children's health and education (see for e.g., Loughran and Zissimopoulos, 2004; Field and Ambrus, 2008; Dahl 2010; Sekhri and Debnath, 2014; Assaad et al., 2017; Chari et al., 2017; Wang and Wang, 2017; Asadullah and Wahhaj, 2019; Sunder, 2019). Of the studies that examine the effect of women's age at marriage on labor market outcomes, Loughran and Zissimopoulos (2004), Dahl (2010) and Wang and Wang (2017) have focused on the United States. Given that gender norms, women's role in society, and socioeconomic status of women in the United States or other developed countries is vastly different from that in developing countries (for instance, the male backlash effect is likely to be much less salient in developed countries compared to developing countries), the studies of Loughran and Zissimopoulos (2004), Dahl (2010) and Wang and Wang (2017), although important, are unlikely to be useful for understanding the link between age at marriage and women's labor market outcomes in developing countries.

To our knowledge, only Assaad et al. (2017) and Sunder (2019) have attempted to examine the causal relationship between age at marriage and women's labor market outcomes in context of developing countries. Our paper complements these studies but differs from them in two key ways. First, in both Assaad et al. (2018) and Sunder (2019), the only labor market outcome considered is labor market participation. We, on the other hand, not only examine the effect of women's age at marriage on labor market participation, but also on other labor market outcomes such as hourly and annual earnings and work days per year. Second, our study context is different from Assaad et al. (2017) and Sunder (2019). While Assaad et al. (2017) and Sunder (2019) focus on African countries, our context of study is India.

This paper is also related to the literature that analyzes the causes of dismal labor market outcomes, especially stagnant labor force participation, of Indian women. Previous studies in this literature have attributed the stagnation of female labor force participation to supply side factors such as rising household incomes, husband's education, increasing returns to home production of educated women, and demand side factors including lack of employment opportunities of women (see Klasen and Pieters (2015) and Afridi et al. (2018) for an excellent overview of this topic). However, so far, the existing literature on female labor force participation has not looked at whether early age at marriage could be held as contributing factor despite economic theory suggesting a negative link between age at marriage and women's labor market outcomes.

The rest of the paper unfolds as follows. In the next section, we discuss the background of our study. In section 3 we discuss the dataset used. Section 4 presents the econometric model and empirical strategy. Results are presented in the section 5. The last section concludes.

2 Study Context and Background

According to the Prohibition of Child Marriage Act (PCMA) 2006, the minimum legal age of marriage in India is 18 years for girls and 21 years for boys with no exceptions.[3] The PCMA 2006 declares child marriage to be a cognizable and non-bailable offence. Despite this, close to 30% of Indian girls are married before their 15th birthday, and almost 1 in 3 child brides worldwide are in India (UNICEF, 2014), .

As per the National Family Health Survey 2015-16 data, child marriage can be seen across India but it is far higher in rural than in urban areas. Further, girls from poorer families and excluded communities–scheduled castes and tribes–are more likely to marry at a younger age. According to UNICEF (2016), the states with the highest prevalence of child marriage (50 per cent and above) are Bihar, Rajasthan, Jharkhand, Uttar Pradesh, West Bengal, Madhya Pradesh, Andhra Pradesh and Karnataka. However, even in states with overall lower prevalence of child marriage, there are often pockets of high prevalence.

[3] http://legislative.gov.in/sites/default/files/A2007-06.pdf

While child marriage can happen to both boys and girls, the practice mostly affects girls. Girls often get married early because of pressure from parents and relatives, poverty, gender norms and lack of alternatives. As noted by Jensen and Thornton (2003) and UNICEF (2016), families may be unwilling to postpone their daughter's marriage due to the high premium placed on female virginity and fears of loss of sexual purity. Limited access to quality education and families' prioritization of boys' rather than girls' education–in part because of limited job opportunities–contribute to perpetuate the practice. Law enforcement to prohibit child marriage is also relatively weak. Limited detailed knowledge on how to apply laws and little understanding of the consequences of the laws, as well as limited trust in institutions enforcing them, undermines the implementation of the PCMA 2006.

Like early marriage, another important issue that concerns women currently in India is their dismal labor market prospects. In 2011, according to the Afridi et al. (2018)'s estimates using NSS data, only 33% of adult Indian women had a job or were actively looking for one. This figure declined to 27% in 2018. As noted in an article in *The Economist* (July 5, 2018), "[Indian] women are less likely to work than they are in any country in the G20, except for Saudi Arabia. They contribute one-sixth of economic output, among the lowest shares in the world and half the global average." The earnings and wages of Indian women who are working are also notoriously low. According to the Global Wage Report 2018-19 published by the International Labour Organization (ILO), hourly wages of women are 34% less than men in India. This gap in wages is highest among 73 countries studied in the report. This is not surprising because women workers are mostly engaged in agricultural sector where the pay is the lowest (as per the 2011-12 NSSO estimates around 62% of women workers are engaged in agriculture; the corresponding figure of male workers is 43%). As Raveendran (2016) notes, women's share in high paying senior and middle management positions in service sector is much smaller than that of men. The place of work of about 34.9% of women workers in 2011–2012, as against 11.4 per cent of men, was own dwelling or adjacent areas. These workers are known as homebased workers and are either self-employed or receiving wages on a piece-rate basis for the amount of work done. As summed up by Raveendran (2016, p. 1), "the deprivation of women in terms of quality of work is three times that of men, and wage rates of women are significantly lower."

In this study, we seek to examine whether the above two phenomena characterizing the Indian economy at present are inter-related. Specifically, we explore whether the poor labor market prospects of women in India could be explained by the high incidence of early marriage. In theory, early marriage hampers labor market prospects of women in multiple ways. First, early marriage causes girls to drop out of school early (Field and Ambrus, 2008; Lloyd and Mensch, 2008; Nguyen and Wodon, 2017). Married girls who live far from school may not be able to continue their education because traveling long distances might simply be infeasible due to lack of transportation facilities or because it represents a risk to be harassed on the way to school (Wodon et al., 2017). Married girls could drop out of school also because education after marriage is not encouraged by their spouses and in-laws since a girl's role of wife often comes with the expectation that she would devote all her time towards taking care of the home and the extended family (Mensch et al., 1998).

The second channel through which women's age at marriage might impact her labor market outcomes is fertility. Early marriage also implies early motherhood which increases the value of women's time at home. This causes younger brides to focus more on home production (specifically, the production of "child services") and reduces their likelihood of participation and productivity in the labor market (Wang and Wang, 2017). In fact, as noted in Wodon et al. (2017), sometimes, the stigma of pregnancy itself keeps away young brides from continuing school which in turn hampers their labor market prospects.

Next, women's age at marriage can affect labor market outcomes through transmission of social norms and gender attitudes. In a carefully designed study conducted in context of Bangladesh, Asadullah and Wahhaj (2019) find clear evidence that the practice of early marriage serves as a conduit for transmission of traditional norms discouraging female employment. In particular, Asadullah and Wahhaj (2019) find that early marriage increases the likelihood of women agreeing with statements like: 'When a woman is engaged in work outside of the home, her children suffer because they are deprived of their mother's attention', 'A woman should not earn more than her husband as this can cause tensions within the household', etc. If this phenomenon is true for India as well, this might be another channel through which early marriage hampers labor market outcomes of Indian women.

Finally, women's age at marriage might influence their labor market outcomes through the channel of dowry. The demand of youthful brides in patriarchal societies is generally very high. As noted by Field and Ambrus (2008), the reasons for this preference include the beliefs that younger brides are (a) more fertile, (b) more likely to lack sexual experience, and (c) easier for husbands and in-laws to control. The high demand of youthful brides could cause early marriage to be associated with lower dowry. If younger brides bring in less assets to their spousal household as dowry, this might negatively influence their intrahousehold decision power (Brown, 2009) and, in turn, hamper their labor market prospects.[4]

3 Data

The data come from the Indian Human Development Survey (IHDS) 2012. IHDS 2012 is a nationally representative multitopic household survey conducted by the National Council for Applied Economic Research (NCAER) in New Delhi and University of Maryland (Desai et al. 2015). The survey was conducted between November 2011 and October 2012, covers 42,152 households located throughout India. The survey covered all the states and union territories of India except Andaman and Nicobar, and Lakshadweep. These two account for less than 0.05 percent of India's population. The data is publicly available from the Data Sharing for Demographic Research program of the Inter-university Consortium for Political and Social Research (ICPSR).[5] The sample was drawn using stratified random sampling.

The IHDS sampled ever-married women above the ages of 15 (one was randomly chosen

[4] See Parsons et al. (2015) for further discussion on how early marriage affects labor force participation of women.

[5] http://www.icpsr.umich.edu/icpsrweb/DSDR/studies/36151

from each surveyed household), who were then administered a separate health and education questionnaire that included questions on marriage and reproductive history, as well as questions on health investments. For our analysis, we restrict ourself to the women whose marital age is not less than 5 years and menarcheal age between 9 and 21 years,[6] have valid information on age, height, caste, family attributes like parental education and number of siblings, and place of residence (rural/urban), leaving us with 37,655 women.

In our analysis, we specifically focus on four labor market outcomes of women: labor market participation, hourly earnings,[7] annual wage earnings, and work days per year. Out of 37,655 women in our sample around 50% have participated in farm, business or worked for wages or salary in last one year (i.e., the year preceding the survey year).[8] Of these women we have valid information on earnings and hours worked for 10,511 of them. The average hourly earnings of the women in our analytical sample is Rs. 18.25, average annual wage earnings is Rs. 24,000, average number of work days per year is 205. The average age of marriage of women is 17.91 years. The average age at menarche is 13.90 years. The average age of women is around 36 years. The mean height is 152 cm. On average, women in our analytic sample have completed 5 years of formal schooling, their fathers have completed 3 years of formal schooling and their mothers have completed less than 2 years of formal schooling. The average number of siblings that the women have is 4. In terms of caste affiliation of the women in our analytical sample, 5% are Brahmins, 23% belong to the general caste category, 41% are members of other backward classes (OBC) and around 29% belong to either schedule caste (SC) or schedule tribe (ST). Finally, 34% of the women in our

[6] The normal menarcheal age is between 10 and 15 years. However, menarcheal age as low as 9 years is not unusual (see for e.g. https://timesofindia.indiatimes.com/city/goa/Girl-talk-Menarche-now-at-8-9-years/articleshow/34169175.cms). Similarly, menarcheal age above 15 years, and in fact, as high as 20-21 years is also not biologically impossible. Delayed puberty may be constitutional or due to pathologic causes (Blondell et al., 1999). Uundernourishment during childhood is, in fact, one major reason for delayed menarche. Also, intense physical activity during childhood may delay menarcheal age. In this context, based on a survey of dancers and atletes, Frisch et al. (1980) and Frisch et al. (1981) note that dancers and athletes who began their training at ages 9 or 10 years still had not menarche at ages 18–20 years.

[7] Hourly earnings include hourly wages, bonus and other in-cash or kind benefits.

[8] Among the working women, around 47% of the women work as the agricultural labourers, 23% work in construction, 6% work as teachers and the rest work in other areas.

sample reside in urban areas, while the rest reside in rural areas. Note, for our analysis we use hourly earnings and annual wage earnings in logs. Table 1 provides descriptive statistics on the analytical sample. Figures 1 and 2 graph the distribution of the age at marriage and age at menarche respectively.

Our estimate of women's labor market participation rate of 50% is significantly higher than what one would estimate based on NSS data. Specifically, Afridi et al. (2018) estimate women's labor market participation to be approximately 33% in 2011 using the NSS data. As suggested by Desai (2017) and Dhanaraj and Mahambare (2019), this difference most likely has to do with how participation in labor market of women has been captured in the two surveys. Specifically, Desai (2017) note:

> "Unlike the NSSO, the IHDS collects data on both income and employment in a single module. Thus, it first asks whether the household owns or cultivates land, then asks about season-wise production, and finally asks who engaged in farm work. Similarly, for wage and salary work, it lists every single paid activity that individuals undertake, regardless of the number of days they work. This allows for a greater capture of fragmented and multiple activities. As a result, IHDS work participation rates for women are higher than the NSS participation rates, but those for men are comparable."

Similar arguments have also been made by Dhanaraj and Mahambare (2019):

> "The advantages of this [IHDS] dataset are: 1) it reduces to some extent the under-reporting of female labour typically associated with censuses and employment surveys in India due to inability to estimate the total work. 2) Women are more often engaged in multiple informal tasks/jobs and the NSSO surveys only capture the main and one or two secondary activities. Thus, she may not have a main activity but perform many small activities (Desai & Jain, 1994) which may not be captured by employment surveys but is captured in IHDS data."

Nevertheless, to ensure that the our results are not an artefact of measurement issues, we use a alternative definitions of women's labor market participation variable in Section 5.3.2 and check the robustness of our baseline results.

4 Empirical strategy

4.1 Econometric model

To examine the impact of women's age at marriage on their labor market outcomes, we begin by estimating the following econometric model:

$$y_i = \alpha + \beta MarriageAge_i + \gamma X_i + \varepsilon_i \tag{1}$$

where y_i denotes a labor market outcome of woman i, $MarriageAge_i$ denotes the woman's age at marriage, X_i denotes the vector of individual and household level controls such as the woman's age, height, family attributes like her father's and mother's years of schooling, number of siblings, place of residence (urban/rural), caste and district fixed effects, and ε_i is the idiosyncratic error term that includes unobserved attributes like ability. Our parameter of interest is the coefficient β which captures the effect of women's age at marriage on their labor market outcomes. If indeed marriage delay leads to better labor market outcomes, β should be positive.[9]

Note, we could have consistently estimated β via OLS estimation and interpreted it as causal effect of age of marriage on labor market outcomes if, conditioning on exogenous characteristics, age at marriage was uncorrelated with unobservable determinants of labor market outcomes (or more formally, $\mathbb{E}[MarriageAge \cdot \varepsilon | X] = 0$). However, such assumption

[9] As most studies in the literature do, we exclude variables such as women's educational attainment and number of children from the set of covariates, as these are potentially endogenous variables that could be influenced by a woman's age at marriage (see for e.g. Wang and Wang, 2017). That is, educational attainment and number of children themselves could be reasons why age at marriage affects a woman's labor market outcomes. Given that we condition on only exogenous variables, the estimated coefficient of β should be interpreted as the *total* effect of women's age at marriage on labor market outcomes.

may be violated due to omitted variables which may affect both the age at marriage of the women and their labor market outcomes. For instance, as noted previously, parents of girls with better labor market prospects (or girls with higher ability) may postpone their daughters' marriages until their earnings capabilities are fully revealed. Also, those parents who care a great deal about their daughter's career could postpone their daughters' marriages in order to let them pursue their careers. Both examples suggest that $\mathbb{E}[MarriageAge \cdot \varepsilon \mid X] \neq 0$ and more likely, $\mathbb{E}[MarriageAge \cdot \varepsilon \mid X] > 0$. As a result, OLS estimates would be biased.[10]

To address this issue, we follow an instrument variable (IV) approach. We use age of menarche as an instrument for women's age at marriage. This instrument is motivated by the observation that has been made by sociologists and anthropologists that parents become extremely anxious to get their daughter married once she has reached menarche, partly to avert any unwanted pregnancies (Caldwell et al., 1983; Srinivas, 1984; Chari et al., 2017). As noted by Field and Ambrus (2008), a significant portion of the variation in timing of menarche is random, rendering it a good instrument for the age at marriage.[11] In what follows, we discuss our IV strategy in detail.

4.2 Instrumental variable strategy

The IV approach involves estimating a two stage model which is specified as follows:

$$MarriageAge_i = \lambda + \delta MenarcheAge_i + \kappa X_i + \eta_i \tag{2}$$

$$y_i = \alpha + \beta MarriageAge_i + \gamma X_i + \varepsilon_i \tag{3}$$

The first stage is given by the equation (2), and equation (3) is the structural equation.

[10] In principal, there might be other potential omitted variables which are not orthogonal to age of marriage of the women and might be correlated with the labor market outcomes considered.

[11] Studies of twins have found that random genetic variation is the single largest source of variations in menarche (see for e.g. Kaprio et al., 1995)

The women's age at marriage, $MarriageAge_i$, is instrumented by $MenarcheAge_i$, their age at menarche, and y_i are the women's and their spouses' labor market outcomes of interest. X_i, as noted above, is a vector of individual and household level controls.

We use a standard two stage estimation procedure (i.e., two stage least squares (TSLS)) and cluster standard errors at the district level.

4.3 Validity of the instrumental variable

In this section, we perform several checks to test the validity of the instrumental variable. The results are presented in Table 2. Results presented in panel A are obtained based on the full analytical sample of women: it is the sample that we would later on use for the regressions pertaining to women's labor market participation. Results presented in panel B, on the other hand, are obtained using the subsample of working women for whom we have valid information on their labor market outcomes: this is the sample which we would later use for the regressions pertaining to hourly earnings, annual wage earnings and work days per year.

We begin by regressing women's age at marriage on age at menarche to examine whether age at menarche predicts age at marriage which is the endogenous regressor. The results are presented in column (1) of both panels. In line with the findings of Field and Ambrus (2008) in context of Bangladesh, and that of Sekhri and Debnath (2014) and Chari et al. (2017) in context of India, we find that age at menarche significantly predicts age at marriage. The value of the coefficient of age at menarche based on the full analytical sample is 0.165, and it is statistically significant at 1% level of significance. The F-Statistic for the regression model is 14.63. The value of the coefficient of age at menarche based on the subsample of working women having valid earnings and days worked information is 0.243. It is statistically significant at 1% level of significance and the F-Statistic for the regression model is 14.87. These results eliminate concerns about 'weak instruments' Additionally, Figure 3 also presents the kernel density estimate of women's age at marriage by menarcheal age groups (early and late menarche)[12] revealing that the distributions of women's age at marriage is positively related to age at menarche.

Next, we examine the potential threats to the validity of this instrument. Medical literature suggests that severe malnutrition in early childhood might result in delayed onset of menarche (Sekhri and Debnath, 2014). Exposure to severe malnutrition could potentially also affect long term health of the women (for e.g. Stathopolu et al. (2003) note that acute malnutrition could result in stunting) and this consequently could affect their labor market prospects. This could undermine our instrument. To address this issue, we include adult height in the regressions in column (2) as a proxy for acute malnutrition in childhood. As noted by Chari et al. (2017), if height is a sufficient statistic for health investments and if undernutrition that affects menarche is also is severe enough to result in stunting, then conditioning on height is likely to eliminate any confounding factor related to health investments that affect both menarche and marriage conditions. We find that inclusion of height as an additional control changes the point estimates and the standard errors only slightly. Even if height is not a sufficient statistic for health, since it is closely related to health (Strauss and Thomas 1989), the fact that controlling for height has very small effects on our results suggests that they are not driven by unobserved health inputs that also affect age at menarche.

As argued by Field and Ambrus (2008), abrupt changes in diet might also affects maturation. Sekhri and Debnath (2014) in this context note that, agriculture and agriculture-related activities, that employ majority of the Indians, are highly weather dependent. Extreme weather conditions (e.g. droughts, floods, etc.) in the women's birth year might lead to loss in household income resulting in transitory but severe malnutrition. Therefore, females born during these extreme weather events may experience delayed age at menarche as they are more likely to be malnourished. We control for this possibility in our first stage regression. In column (3), we add age of the women to account for extreme weather events at the time

[12] The early menarche group consists of those women who attained menarche at the age of 14 or earlier. The late menarche group consists of those women who attained menarche after the age of 14.

of birth. The point estimates and standard errors are similar across columns (1) and (3) in both Panels A as well as B. We condition all subsequent results on adult height and women's age.

As noted by Asadullah and Wahhaj (2019), age at menarche as an IV could be endogenous through correlation with omitted family characteristics. A prime example of such omitted family characteristics is socioeconomic status (SES) of family. Women who come from low SES families are more likely to be involved in strenuous physical labor during early childhood (Sekhri and Debnath, 2014). This can lead to a delay in menarche due to adverse effects of hard physical labor on health (Pellerin-Massicotte et al., 1997). Thus women who end up marrying late may also be less healthy, and this could have a direct effect on her labor market prospects. To address this concern we include controls for women's father's and mother's educational attainment (i.e., years of schooling) as well as the number of siblings of the women and reestimate the first stage equation. Additionally, in this regression we also include controls for women's caste affiliation. We believe that these family characteristics are likely to serve as good proxies for socioeconomic status (SES) of women's natal family. As evident from the results reported in Column (4), the inclusion of the women's natal family characteristics as additional controls does not change the point estimates of the coefficient of age at menarche significantly.[13]

Age at menarche might also be potentially endogenous due to geographical factors such as temperature, rainfall, altitude, etc. (Field and Ambrus 2008; Chari et al. 2017). To address this issue, we control for place of residence (whether the household resides in an urban or a rural locality) and use district fixed effects to account for spatial variation in exposure to environmental factors that affect menarche. Note, we are able to control for district of residence of the married woman, and not her natal district since we do not have

[13] Asadullah and Wahhaj (2019) address the issue of omitted family characteristics by using sisters fixed effects, and exploit only the variation in age of menarche between sisters. To the extent that sisters are raised within the same household by the same parents, this approach allows them to abstract away from variation in the age of menarche that is due to common environmental and socioeconomic factors. Unfortunately, we could employ this strategy due to data limitations.

any information about the location of her natal family. This, however, is not likely to be a problem because in India most marriages occur within the same district, so the district of residence of the married woman is also likely also her natal district (Fulford 2015). The results of the specification that include geographic controls is presented in Column (5). The coefficient of age at menarche is still highly statistically significant.

Another concern is measurement error in the age at menarche (Chari et al., 2017; Asadullah and Wahhaj, 2019). While this is possible since it was self-reported by respondents at the time of the survey, Garg et al. (2001) and Sharma et al. (2006) note that menarche is a major event for girls in India, and girls of both low and high caste report knowing little or nothing about menstruation before it began, but afterwards learning of taboos about eating and mobility during menstrual periods. These changes in lifestyle imply that respondents are likely to recall its timing with fair degree of accuracy (Chari et al., 2017).[14] Furthermore, the distribution of reported age at menarche (Figure 2) does not show any heaping at key ages (e.g. school leaving ages) that might be suggestive of significant recall error.

The final concern that we need to address is whether our instrument is exogenous given that we are not controlling for education in the second stage equation. One might argue that a woman's educational attainment as measured by her years of schooling, is correlated with her age at menarche. More specifically, menarche itself might be a barrier to schooling (as often cited in the popular media). If this is the case, then leaving out education from the set of control variables will violate the condition that $\mathbb{E}[MenarcheAge \cdot \varepsilon |\ X] = 0$, and the IV regressions will not yield consistent estimates of the parameters of interest.

While this is possible, Field and Ambrus (2008) in their seminal paper provide robust evidence that menarcheal age has no direct impact on women's schooling using data from Bangladesh. Oster and Thornton (2011) although document a statistically significant effect of menstruation on school attendance for girls in Nepal, this effect is extraordinarily small.

[14] Ellis (2004, 921) based on a survey also note, "both adolescent girls and adult women are generally willing and able to report accurately on their ages at menarche...and retrospective reports may be more reliable than those obtained during puberty".

Specifically, they estimate that girls miss a total of only 0.4 days in a 180 day school year (although 47% of the girls in their study reported missing some school due to menstruation in the past year). Further, Oster and Thornton (2011) show that improved sanitary technology has no effect on reducing this small gap: girls who randomly received sanitary products were no less likely to miss school during their period. Grant et al. (2013) conduct a study in Malawi to examine the individual and the school level factors associated with menstruation-related school absenteeism. In line with the findings of Field and Ambrus (2008) and Oster and Thornton (2011), they find no evidence that menstrual periods account for female absenteeism. Thus, even though it is often believed that menstruation causes girls to be absent from school, these findings indicate that in reality it is unlikely to be the case.

Nevertheless, to address the concern that our instrument might potentially be endogenous due to omission of schooling from our model, we do the following. First, we plot the average years of schooling of women by different menarcheal age in Figure 4. We find no evidence of an upward trend in the relationship between schooling and age at menarche. Second, we present the kernel density estimate of women's years of schooling by terciles of menarcheal age in Figure 5. The figure reveals that the population distributions, and not just averages, are remarkably similar across all subsamples. This is not what we would have expected to find if menarcheal age was correlated with years of schooling. This suggests that not controlling for educational attainment of women is unlikely to confound our analysis.[15]

[15] Note, Sekhri and Debnath (2014) and Chari et al. (2017) also implicitly assume that age of menarche is not correlated with women's education. Both the papers investigate the impact of marital age of the mother on child health and education outcomes. Marital age is instrumented by menarcheal age, but mother's education is not controlled for. Given that mother's education is conjectured to a determinant of child outcomes, mother's education becomes of the part of the error term in the second stage regression, which must be assumed to be uncorrelated to menarcheal age, for their second stage parameter estimates to be consistent.

5 Results

5.1 OLS results

The OLS estimates of the effect of women's age at marriage on their labor market outcomes are presented in Table 3. While these estimates are not causal, nevertheless they are likely to serve as useful benchmarks with which we would be able to compare our IV estimates, in turn, allowing us to distinguish causality from correlation due to selection into marriage.

Examining the results of Table 3, we first find that women's age at marriage is negatively associated with labor market participation (column 1) but positively associated with all other labor market outcomes namely hourly earnings (column 2), annual wage earnings (column 3), and work days per year (column 4). More specifically, our results indicate that a year of delayed marriage decreases the probability of women's labor force participation by 0.7%, women's labor market participation probability by women's hourly earnings by 2%, annual wage earnings by 3%, and work days per year by roughly 2. All the estimated coefficients are statistically significant at 1% level of significance. These results imply that marital delay leads women to participate less in the labor market, but it is beneficial for those women who do participate.

To examine whether the observed relationship between women's age at marriage and their labor market outcomes is causal, we next turn our attention to the IV results.

5.2 IV results

The IV results are presented in Table 4. Our findings are as follows. First, in sharp contrast to the OLS results, we find that IV estimates of the coefficients of women's age at marriage are not statistically significant in any of the regressions. Second, in terms of the size of the estimated coefficients, the IV estimates of the coefficients of women's age at marriage are drastically smaller in absolute magnitude compared to the OLS estimates of the coefficients in all the regressions. For example, the effect of a one year delay in women's age at marriage on their hourly earnings is now only 0.5%. The corresponding OLS figure was 2%. Similarly, the effect of a year delay in women's annual wage earnings is now 1%, which according to the OLS estimates was 3%. Third, in terms of the direction of the effect, the IV estimates of the coefficients of age at marriage in the regressions with women's labor market participation, hourly earnings and annual earnings as dependent variables retain the same sign as that of the OLS estimates; however, the IV estimate of the coefficient of age at marriage in the regression with women's work days per year as the dependent variable is of the opposite sign as compared to the sign of the OLS estimate.

Overall, thus the IV results indicate that a delay in marriage of women by a year has no statistically or economically significant causal impact on their labor market outcomes. The OLS results were most likely arising due to the influence of the omitted variables.

5.3 Robustness Checks

5.3.1 Comparison of OLS and IV based on compliers: selection vs. causal mechanisms

The comparison of our OLS and IV estimates indicates that our results are not consistent with the causal hypotheses. However, as noted by Wang and Wang (2017), the comparison of full-sample OLS and IV estimates may not necessarily be a fair evaluation of these hypotheses. As pointed out in Imbens and Angrist (1994), the IV estimates capture only the effect for a subgroup of individuals who are likely impacted by the IV (i.e., the compliers). The IV estimates for the complier sample might actually be higher than the IV estimates for the full sample.

In the present study, the set of compliers will not consist of women who were married before they attained maturation. For these women, age at menarche would not have affected their marital age. Consequently, we assume that our complier subsample excludes women who were married before menarche. We re-estimate the OLS and IV models for these women. The OLS and IV results for the potential compliers are reported in Tables 5 and 6 respectively.

We continue to find that the OLS estimates are larger than the IV estimates. Moreover, as before, the OLS estimates are statistically significant whereas the IV estimates are not. This again indicates that there does not exist significant causal relationship between age at marriage and women's labor market outcomes.[16]

5.3.2 Alternative definitions of labor market participation

Our baseline definition of women's labor market participation considers a woman to be participating the labor market if she has participated across all types of work in last one year. Under this definition, around 50% women seem to be participating in the labor market. As noted previously, this participation rate is significantly higher than the participation rate estimated based on the NSS data. To assess whether our results for labor market participation are an artefact of how we define women's labor market participation we use two alternative definitions and check if our baseline results hold.

Our first alternative definition follows Chatterjee et al. (2018)'s definition of women's labor market participation. According to this definition a woman is considered to be participating in the labor market only if she has worked for *at least* 240 hours in the previous year across all types of work. We did not have such a cut-off for our baseline definition according to which all women who were participating in any work for any length of time in the preceding year were considered to be participating in the labor market. Under this alternate definition, women's labor market participation rate falls from around 50% to close to 42%. Our second alternative definition is based on women's report on whether they have *ever* participated in the labor market. Under this definition, a woman is considered to have participated in the labor market, if she has *ever* worked (instead of just considering the last one year as the relevant period).

[16] It might be that the complier set excludes not only women who were married before maturation but also those who were married *much* after maturation (say, those who were married after teenage or after early adulthood). Consequently, for checking sensitivity of our results, we assume that our complier subsample excludes women who were married before menarche or were married ten years after menarche. The results —reported in the Tables B1 and B2 in the Appendix—indicate no substantial difference in the our findings with respect to that obtained based on the baseline complier sample.

We use these alternative indicator variables of women's labor market participation and re-estimate the baseline regression model to examine the causal effect of women's age at marriage on labor market participation. The results are reported in Appendix Table B3. Thankfully, our baseline results remain qualitatively unchanged. This is reassuring and suggests that our baseline results for women's labor market participation are unlikely to be arising due to measurement issues.

5.3.3 Unaccounted regional heterogeneity

India is a culturally diverse country. Gender norms regarding female employment and perceptions about role of women in society significantly vary across regions in India. While in some regions (e.g., northern states) gender norms discourage women's participation in the work force, in other regions (e.g., southern states), gender norms are supportive of women's outside employment. As noted previously, Asadullah and Wahhaj (2019) in a recent study present clear evidence that the practice of early marriage serves as a conduit for transmission of traditional norms discouraging female employment in Bangladesh. If the findings of this study are valid in context of India, it might be the case that in regions where gender norms discourage (are supportive of) women's work, early marriage might be negatively (positively) affecting female labor force participation and productivity. In other words, there might be some regions where age at marriage is negatively related to women's labor market outcomes, and there might be some regions where age at marriage is positively related to women's labor market outcomes. Our baseline analysis might simply not have picked up these heterogenous effects since we carried out our analysis based on the pooled sample.

To explore this issue, we carry out a subsample analysis for different regions of India: North, South, Central (which includes West and East Indian states as well), and North East. Specifically we focus on five different labor market outcomes: labor market participation in the preceding year, 'ever' labor market participation, hourly earnings, annual wage earnings and work days per year. For each of these labor market outcomes we estimate how it is impacted by women's age at marriage across the four different regions. Results are reported in Appendix Table B4.[17]

We find that 'ever' labor market participation, hourly earnings, annual wage earnings and work days per year are not impacted by women's age at marriage in any of the regions. For labor market participation in the preceding year as well, age at marriage has no effect on it for North, Central and North East regions. However, we do document a statistically significant negative effect of age at marriage on labor market participation in the preceding year when we use the South sample (i.e., we reject the null hypothesis that age at marriage has no impact of women's labor market participation in the preceding year for the South sample). However, since out of the 20 coefficients of women's age at marriage estimated (5 outcomes $\times$ 4 regions), only 1 appear to be statistically significant, by and large, we can claim that, overall, there is no substantial heterogeneity in the relationship between Indian women's age at marriage and their labor market outcomes (and that our baseline results are unlikely to be the outcome of pooling of data across regions with opposite gender norms). This is because when carrying out simultaneous testing of several hypotheses, as noted by Romano et al. (2010, p. 94) in an influential paper, "there typically results a large probability that some of the true hypotheses will get rejected by chance alone." Romano et al. (2010, p. 94) illustrate their point with the following example: "Take the case of $S = 100$ hypotheses being tested at the same time, all of them being true, with the size and level of each test exactly equal to α. For $\alpha = 0.05$, one expects five true hypotheses to be rejected. Furthermore, if all tests are mutually independent, then the probability that at least one true null hypothesis will be rejected is given by $1 - 0.95^{100} = 0.994$."

5.3.4 Issue of selection into labor force

To this point, for analyzing the effect of age at marriage on women's hourly earnings, annual wage earnings and work days, we could consider only women who are working. Given that

[17] See note below Table B4 for a complete list of states included in the categories, North, South, Central and North-East.

non-working women might systematically differ from working individuals, our analysis could be biased. To address this issue, in the appendix, we examine whether the observed results are robust to addressing selection into labor force. Our approach closely follows that used by Wang and Wang (2017) (see Appendix A for details). Thankfully, all of our baseline results continue to hold (see Appendix Tables B5 and B6). Not only do we find similar patterns in our estimates; we generally find estimates to be remarkably similar in magnitudes as well.

5.4 Discussion

Our results are surprising since previous literature on marriage markets in India asserts that women's age at marriage leads to an increase in education and dowry, and a decrease in fertility (see Chowdhury, 2009; Sekhri and Debnath, 2014; Chari et al., 2017). As such, it is expected that a delay in age at marriage is likely to improve women's labor market outcomes in India. One reason, of course, why we do not find the anticipated effect could be that for our sample specifically, none of the pathways mediating the relationship between women's age at marriage and labor market outcomes are functional. To examine this issue, we examine the following pathways: (1) education, (2) fertility, and (3) gender norms. We could not test the effect of age at marriage on dowry due to unavailability of required data on dowry.

Testing the impact of women's age at marriage on education and fertility is relatively straight forward: to test the first (second) channel, we use women's completed years of schooling (number of children born) as the outcome variable. However, testing whether women's age at marriage impacts transmission of gender norms regarding female employment is not easy since measuring gender norms related to work and how much women adhere to such norms is difficult. In this paper, we address this issue by using three alternative outcome variables capturing gender norms related to working outside home.

First, using questions on whether women respondents ask for permission from their husband or a senior family member to go to the local health center, to visit friends or neighbors, to the grocery shop, and to short distance by train or bus, following Desai and Banerji's (2008) strategy, we construct a mobility index. We consider that a woman has to seek permission if she has to inform or explicitly take permission from their husband or a senior family member. Our mobility index ranges from 0 to 4 based on the responses of these four questions. Respondents who need permission for all four visits get an index score 4 and they are considered as least mobile and those who do not need any permission for these visits get an index score zero, and they are considered as most mobile. According to traditional gender norms, women are expected to seek permission from the spouse or older family members before stepping out of the household for various works (Bloom et al., 2001). As such, higher values of this index reflects prevalence stronger gender norms discouraging female employment.

Second, we create a binary outcome variable based on women's response to the following question focus on another question: "do you practice ghungat / burkha/ purdah/ pallu?" Practice of this exercise is considered as the norm of female seclusion and tends to restrict women's mobility.

Third, we construct another index measuring the autonomy of women in various decisions such as what to cook on daily basis, whether to buy an expensive item such as TV or fridge, number of children to have, what to do if the respondent falls sick, whether to buy land or property, amount of money to spend on a social function such as marriage, what to do if a child falls sick, whom your child should marry. Woman respondent answers whether the following household members decide or not: the respondent herself, her husband, senior male, senior female, others or no one. In case of more than one decision maker, she decides who is the primary decision maker for every decision. Our decision making index ranges from 0 to 8 based on the responses of these eight decisions. Respondents who are (are not) the sole or primary decision makers for all (any) the decisions get an index score of 8 (0) and they are considered to have full (no) autonomy in these eight decisions.

Once we have all the outcome variables for testing all three channels, we carry out a set of regressions. In each regression, we regress an outcome variable on women's age at marriage and the set of control variables discussed previously. As before, we instrument women's age at marriage by age at menarche. The results are presented in Table 7.

As evident, for our sample, we find clear evidence that a delay in marriage leads to an increase in educational attainment and decline in fertility, although we fail to find any significant labor market effects of delay in women's age at marriage through transmission of norms supportive of outside employment. However, since we find strong evidence in favor of two out of the four channels being functional, delay in age at marriage should have resulted in beneficial labor market outcomes for Indian women. Why does not a delay in age at marriage then improve women's labor market outcomes in our study?

We offer a possible explanation based on the theory of male backlash. As per the sociological theory of male backlash, education might make women more vulnerable to violence from their spouses and in-laws in a patriarchal society (Macmillan and Gartner, 1999). This is because women's education or empowerment acts as a threat to men's position of superiority in a relationship. As noted by Macmillan and Gartner (1999), women's education or decision power might make things worse since when gender roles and power relations are redefined, men resort to violence to reinstate a culturally prescribed norm of male dominance and female dependence. In India, in particular, the prevalence of male backlash effect has been documented in some sociological studies. Weitzman (2014), for example, using data from the National Family Health Survey (NFHS) 2005-06 finds that education makes married India women more vulnerable to male backlash. Also, in a recent report by the philanthropic organization Dasra (2019), similar findings have been noted.

Since in our sample older brides are more educated or more empowered than the younger brides (see Table 7 and Figure 6), and also the gap between spouses' and older brides' years of schooling is also substantially lower compared to the gap between spouses' and younger brides' years of schooling (see Figure 7), it might be the case that the older brides face more violence from their spouses and in-laws compared to younger brides. This could be negatively impacting the labor market participation and labor market productivity of the older brides. This, in turn, could be offsetting the positive effects of marriage delay on women's labor market prospects through better education, higher dowry and lower fertility.[18]

6 Conclusion

In this paper, we examine the relationship between women's age at marriage and their labor market outcomes using nationally representative household data from India. We use an instrumental variables-based empirical strategy that utilizes variation in age at menarche to obtain exogenous variation in women's age at marriage. We find that there is no significant causal effect of women's age at marriage on their labor market outcomes. This is despite a delay in marriage being clearly associated with more education, lower fertility and (possibly) higher dowry for Indian women. We argue that this might be because, as compared to younger brides, older brides face more backlash from their partners and in-laws and are given less freedom to work.

Our findings suggest that, in contrast to the conventional wisdom, marital delay policies and legislations to prevent child marriages, by themselves, are unlikely to lead to an improvement in Indian women's labor market prospects. Such policies might, however, benefit women in the labor market if they are complemented by interventions that reduces the male backlash effect.

[18] One can argue that that for the set of women whose marriage age is affected by their menarche age, an increase in their marriage age leads to only a modest gain in education; say, marriage delay causes them to move from primary to secondary school. This gain in education might not be sufficient to generate better labor market outcomes. We, however, feel this is unlikely to be valid explanation for our results for two reasons. First, labor market returns and consumption returns to education of women, even at relatively low educational levels, are quite high in India, and in fact they are higher than that of the men (Fulford, 2014; Mitra, 2019). Thus, even if an increase in age in marriage leads to modest gains in education, this should get translated into better labor market outcomes. Second, even if for the sake of argument we assume that for the compliers, an increase in marriage age leads to a modest gain in education which is insufficient to get translated into better labor market outcomes, for the overall impact to be zero, the reduction in fertility and increase in dowry should also be close to zero (since the total effect of age at marriage on women's labor market outcomes is driven by not only education, but also fertility and dowry). We are not sure why would that be the case.

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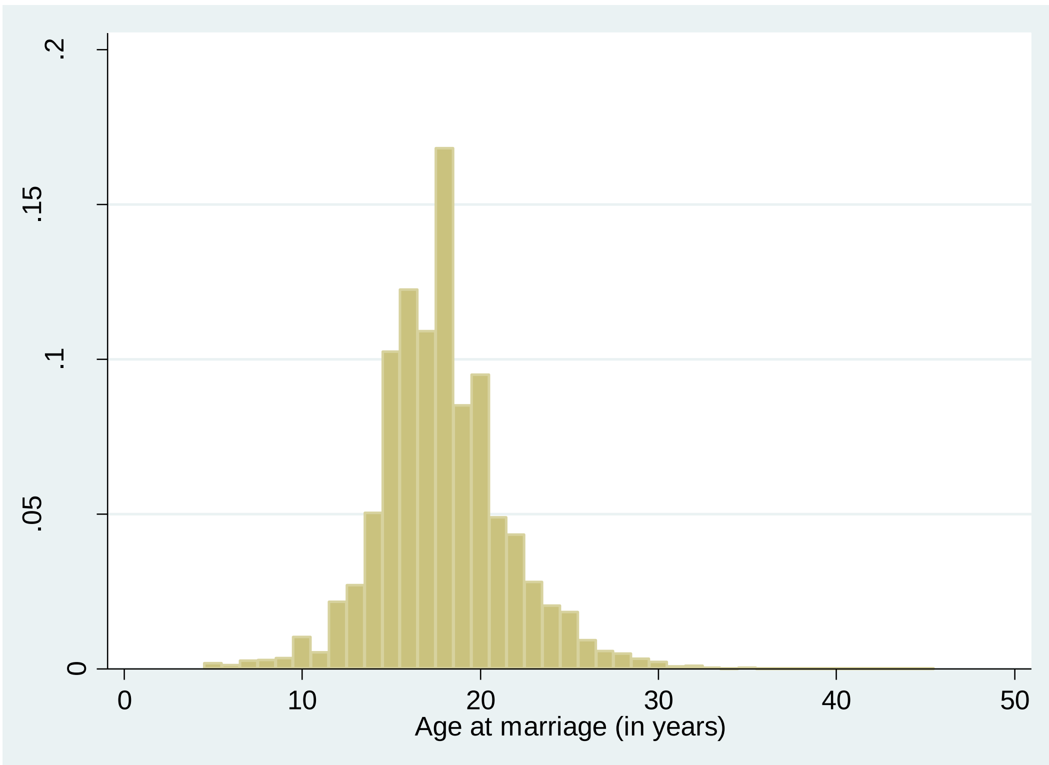

Figure 1. Distribution of women's age at marriage

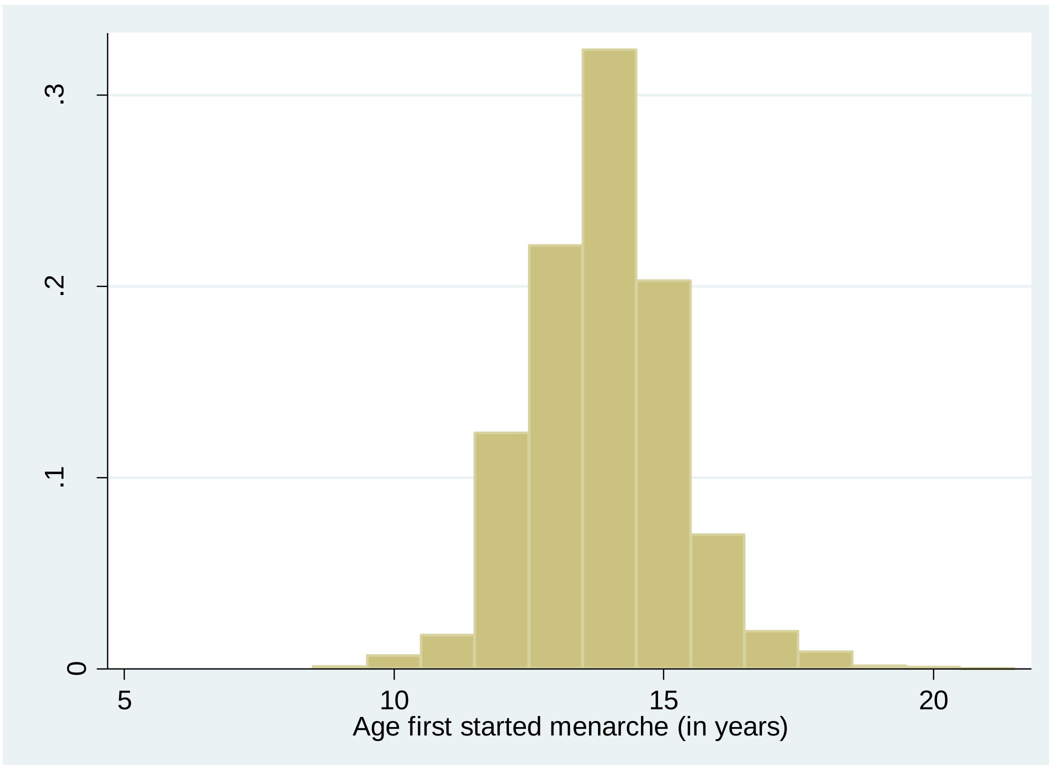

Figure 2. Distribution of age at menarche

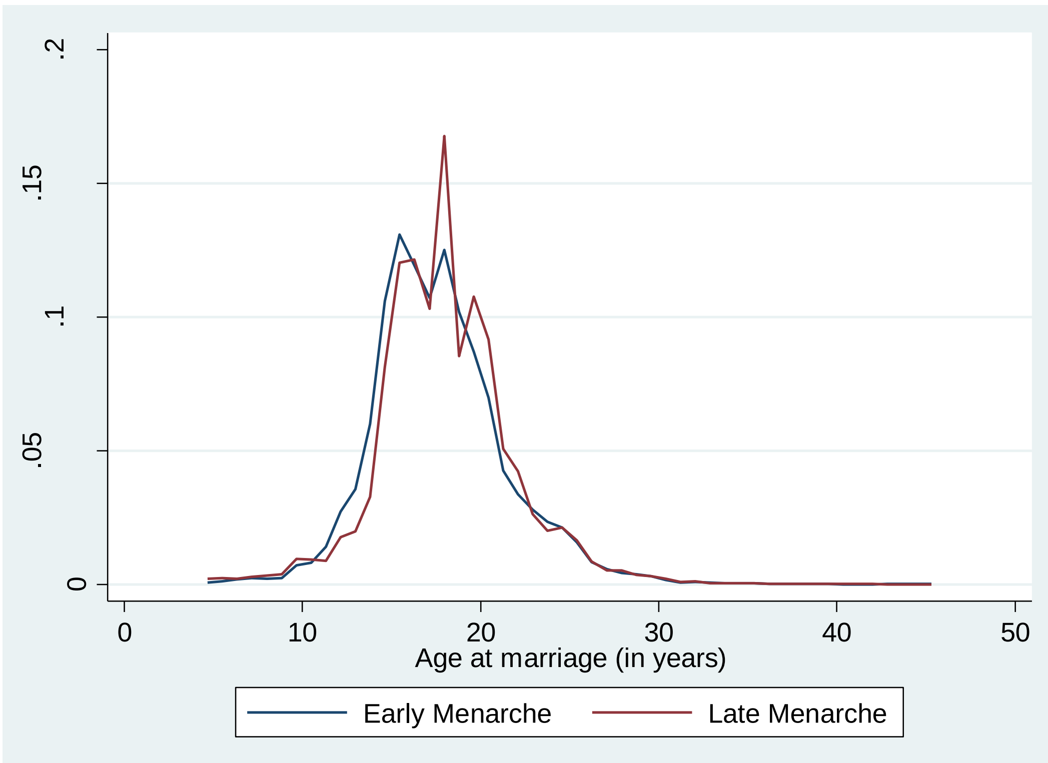

Figure 3. Distribution of women's age at marriage by age at menarche

Notes: Early Menarche implies menarche age ≤ 14 years; late menarche implies menarche age > 14 years

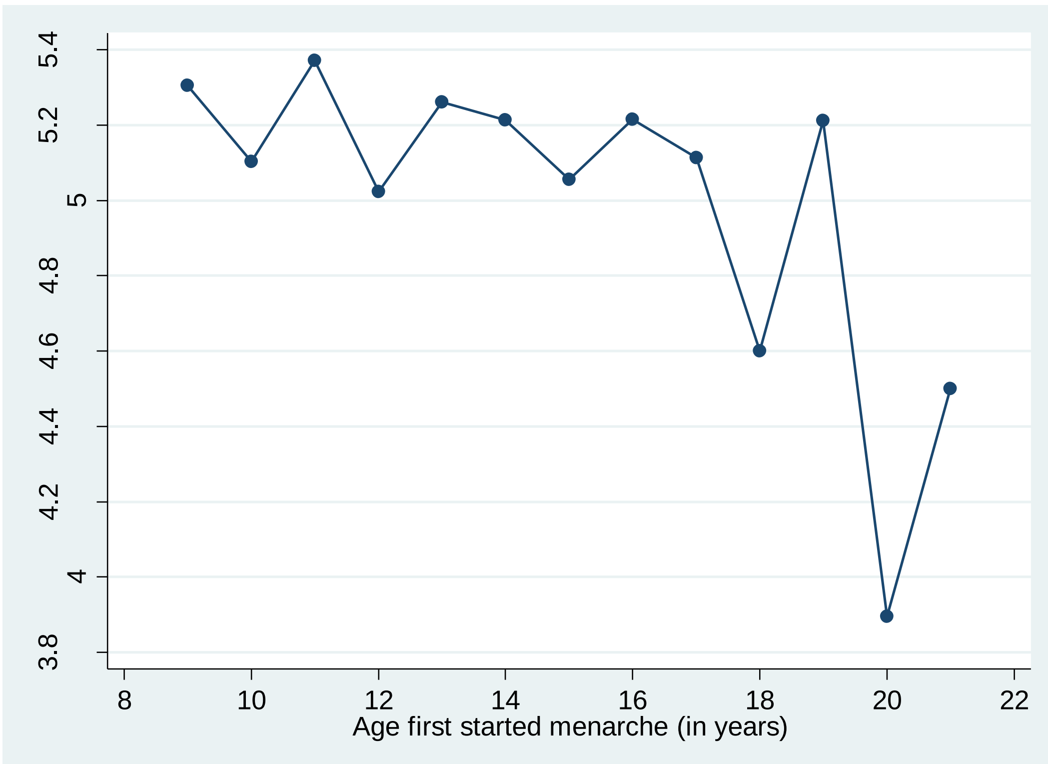

Figure 4. Relationship between women's average years of schooling and age at menarche

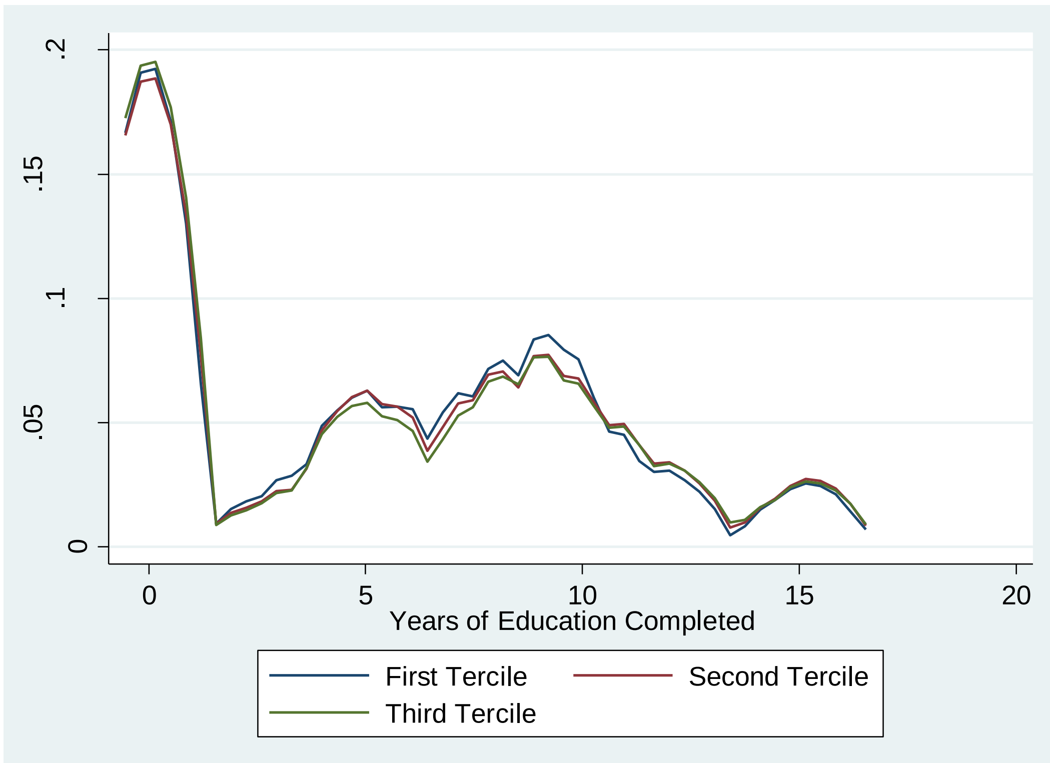

Figure 5. Kernel density estimates of women's years of schooling by terciles of age at menarche

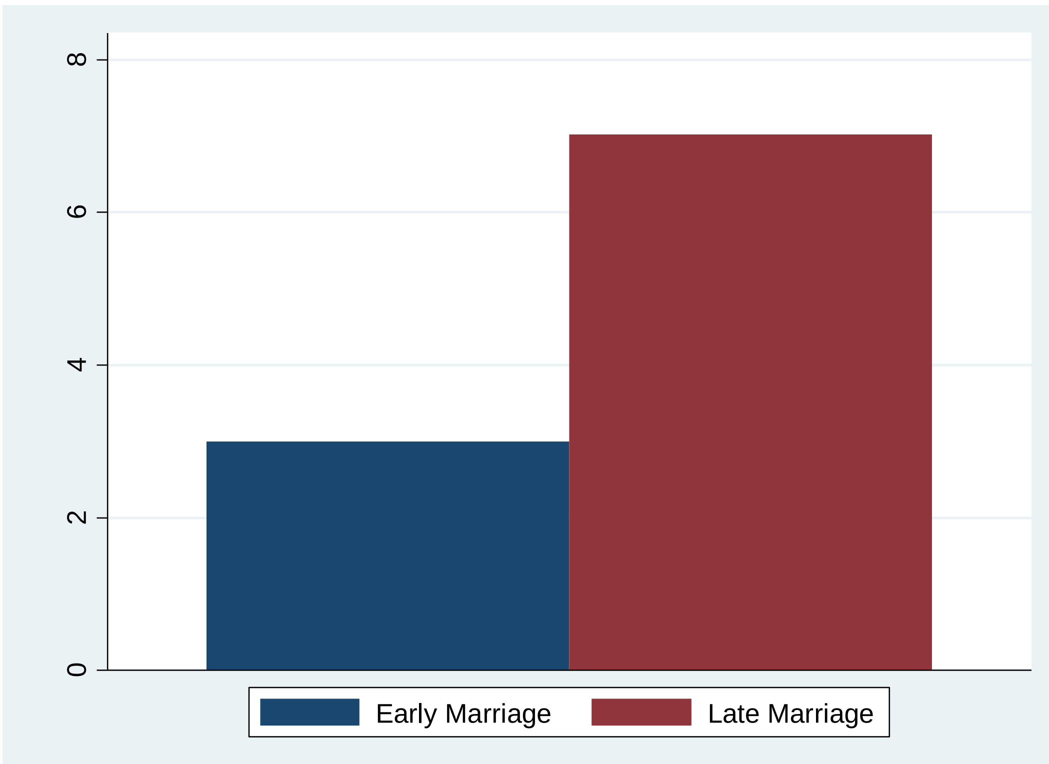

Figure 6. Women's average years of schooling by age at marriage

Notes: Early Marriage implies marriage age < 18 years; late marriage implies marriage ≥ 18 years

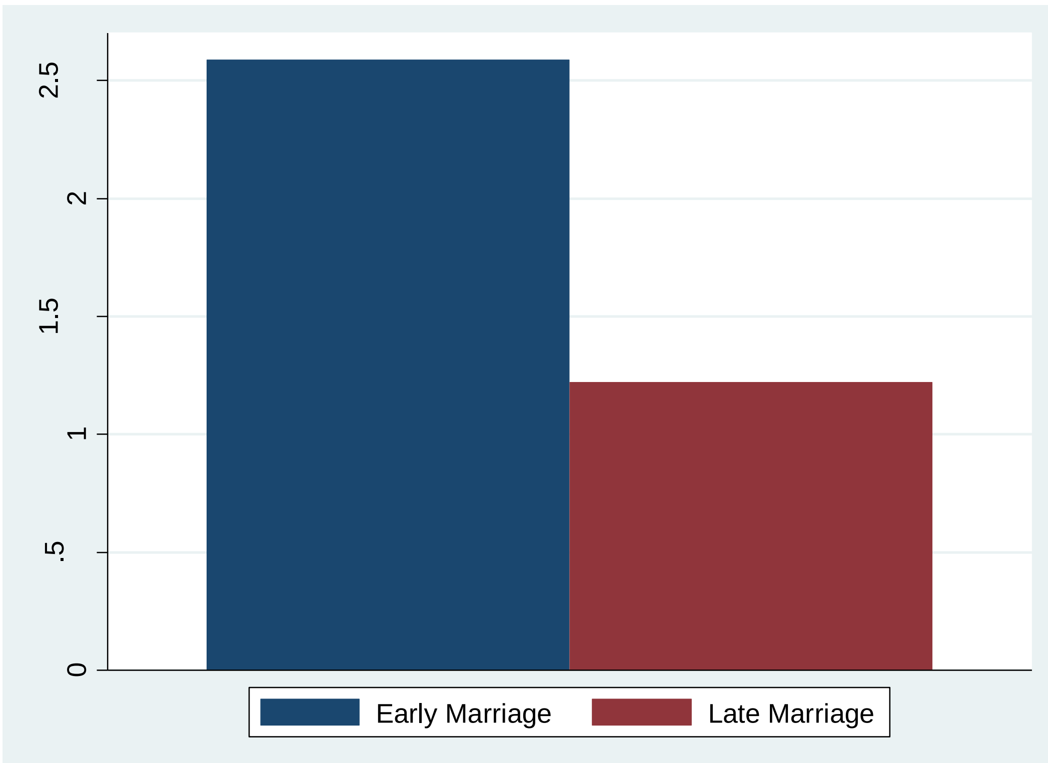

Figure 7. Difference between husbands' and women's years of schooling by women's age at marriage

Notes: Early Marriage implies marriage age < 18 years; late marriage implies marriage ≥ 18 years

Table 1. Summary Statistics

	N	Mean	SD
Women's labor market outcomes			
Participation in labor market (= 1 if participated in farm, business or worked for wages or salary in last one year; = 0 otherwise)	37655	0.50	0.50
Hourly earnings (in Rs.)	10511	18.25	24.40
Annual wage earnings (in Rs.)	10511	23977.56	50282.33
Work days per year	10511	205.29	103.85
Women's demographics			
Age at marriage	37655	17.91	3.67
Age at menarche	37655	13.90	1.39
Age	37655	36.35	9.84
Height (in cm)	37655	152.02	6.63
Years of schooling	36755	5.16	4.92
Father's years of schooling	37655	3.43	4.50
Mother's years of schooling	37655	1.48	3.04
Number of Siblings	37655	3.85	1.99
Brahmin (= 1 if yes; = 0 otherwise)	37655	0.05	0.22
General caste except Brahmin (= 1 if yes; = 0 otherwise)	37655	0.23	0.42
Other Backward Classes (OBC) (= 1 if yes; = 0 otherwise)	37655	0.41	0.49
Scheduled Caste (SC) (= 1 if yes; = 0 otherwise)	37655	0.21	0.41
Scheduled Tribe (ST) (= 1 if yes; = 0 otherwise)	37655	0.08	0.28
Other Castes (= 1 if yes; = 0 otherwise)	37655	0.01	0.11
Place of Residence (=1 if Urban)	37655	0.34	0.47

Table 2. OLS estimates of the effect of age at menarche on age at marriage

Panel A: Full Analytic Sample					
	[1]	[2]	[3]	[4]	[5]
Age at Menarche	0.165***	0.148***	0.152***	0.161***	0.369***
	(0.043)	(0.043)	(0.042)	(0.034)	(0.021)
F-statistic	14.63	44.99	53.82	129.38	300.53
R^2	0.004	0.013	0.019	0.139	0.317
Observations	37,655	37,655	37,655	37,655	37,655
Panel B: Working Women Sample					
	[1]	[2]	[3]	[4]	[5]
Age at Menarche	0.243***	0.227***	0.230***	0.215***	0.446***
	(0.063)	(0.062)	(0.063)	(0.052)	(0.039)
F-statistic	14.87	30.05	21.55	84.00	130.00
R^2	0.008	0.017	0.019	0.149	0.361
Observations	10,511	10,511	10,511	10,511	10,511

Notes: Estimation via OLS. The outcome variable is women's age at marriage. Regressions reported in columns (1) of Panels A and B, do not include any controls. In column (2) regressions we control for women's height. In column (3) regressions the control variables are women's height and age. In column (4), controls include women's age, height, caste affiliation, height, father's years of schooling, mother's years of schooling, and number of siblings. In column (5), we include district fixed effects in addition to all controls used. For regressions reported in columns (2) through (5) in Panel B, we also include spousal age as an additional control. Standard errors reported in the parentheses are clustered at the district level. ***p < 0.01, **p <0.05, *p < 0.1.

Table 3. OLS estimates of the effect of age at marriage on labor market outcomes

	[1]	[2]	[3]	[4]
	Participation	Hourly Earnings	Annual Wage Earnings	Work Days Per Year
Age at Marriage	-0.007***	0.017***	0.033***	1.655***
	(0.001)	(0.002)	(0.004)	(0.352)
R^2	0.285	0.350	0.436	0.278
Observations	37,655	10,511	10,511	10,511

Notes: Estimation via OLS. All regressions control for women's age, caste affiliation, height, father's years of schooling, mother's years of schooling, number of siblings, and district fixed effects. Standard errors reported in the parentheses are clustered at the district level. ***p < 0.01, **p <0.05, *p < 0.1.

Table 4. IV estimates of the effect of age at marriage on women's labor market outcomes

	[1]	[2]	[3]	[4]
	Participation	Hourly Earnings	Annual Wage Earnings	Work Days Per Year
Age at Marriage	-0.002	0.005	0.015	-0.288
	(0.006)	(0.010)	(0.020)	(1.677)
R^2	0.284	0.347	0.434	0.275
First Stage F-statistic	300.53	130.00	130.00	130.00
	[p=0.000]	[p=0.000]	[p=0.000]	[p=0.000]
Kleibergen Paap rK LM statistic	124.13	65.07	65.07	65.07
	[p=0.000]	[p=0.000]	[p=0.000]	[p=0.000]
Observations	37655	10511	10511	10511

Notes: Estimation via TSLS. All regressions control for women's age, caste affiliation, height, father's years of schooling, mother's years of schooling, number of siblings, and district fixed effects. Standard errors reported in the parentheses are clustered at the district level. ***p < 0.01, **p <0.05, *p < 0.1.

Table 5. OLS estimates of the effect of age at marriage on labor market outcomes, Complier Subsample

	[1]	[2]	[3]	[4]
	Participation	Hourly Earnings	Annual Wage Earnings	Work Days Per Year
Age at Marriage	-0.005***	0.023***	0.040***	2.070***
	(0.001)	(0.003)	(0.005)	(0.415)
R^2	0.278	0.364	0.438	0.283
Observations	34634	9362	9362	9362

Notes: Estimation via OLS. All regressions control for women's age, caste affiliation, height, father's years of schooling, mother's years of schooling, number of siblings, and district fixed effects. Standard errors reported in the parentheses are clustered at the district level. ***p < 0.01, **p <0.05, *p < 0.1.

Table 6. IV estimates of the effect of age at marriage on labor market outcomes, Complier subsample

	[1]	[2]	[3]	[3]
	Participation	Hourly Earnings	Annual Wage Earnings	Work Days Per Year
Age at Marriage	-0.003	0.005	0.021	0.072
	(0.005)	(0.009)	(0.017)	(1.429)
R^2	0.278	0.358	0.436	0.281
First stage F statistic	542.17	267.65	267.65	267.65
	[p=0.000]	[p=0.000]	[p=0.000]	[p=0.000]
Kleibergen Paap rK LM statistic	156.35	91.42	91.42	91.42
	[p=0.000]	[p=0.000]	[p=0.000]	[p=0.000]
Observations	34634	9362	9362	9362

Notes: Estimation via TSLS. All regressions control for women's age, caste affiliation, height, father's years of schooling, mother's years of schooling, number of siblings, and district fixed effects. Standard errors reported in the parentheses are clustered at the district level. ***p < 0.01, **p <0.05, *p < 0.1.

Table 7. IV estimates of the effect of age at marriage on women's education, fertility, and prevalence of gender norms related to working outside home

Panel A: Sample for Labor Force Participation

	Years of education		**Number of children born**		**Mobility Index**		**Practice of using ghungat/burkha/purdah/ pallu**		**Decision Making Index**	
	Full Sample	Complier Subsample	Full Sample	Complier Subsample	Full Sample	Complier Subsample	Full Sample	Complier Subsample	Full Sample	Complier Subsample
	[1]	[2]	[3]	[4]	[5]	[6]	[7]	[8]	[9]	[10]
Age at Marriage	0.320***	0.344***	-0.135***	-0.136***	0.004	0.001	-0.001	-0.005	0.017	0.015
	(0.046)	(0.039)	(0.016)	(0.013)	(0.023)	(0.020)	(0.007)	(0.006)	(0.034)	(0.028)
R^2	0.56	0.549	0.445	0.435	0.205	0.21	0.505	0.503	0.143	0.148
First Stage F-statistic	300.46	542.17	303.90	549.8	278.47	472.25	296.96	533.76	278.71	503.37
	[p=0.000]	[p=0.000]	[p=0.000]	[p=0.000]	[p=0.000]	[p=0.000]	[p=0.000]	[p=0.000]	[p=0.000]	[p=0.000]
Kleibergen Paap rK LM statistic	124.11	156.35	125.05	157.29	116.62	142.27	123.49	155.66	121.31	154.7
	[p=0.000]	[p=0.000]	[p=0.000]	[p=0.000]	[p=0.000]	[p=0.000]	[p=0.000]	[p=0.000]	[p=0.000]	[p=0.000]
Observations	37654	34633	37443	34432	29,979	27,685	37,628	34,607	33,333	30,649

Panel B: Sample for Labor Market Outcomes

	Years of education		**Number of children born**		**Mobility Index**		**Practice of using ghungat/burkha/purdah/ pallu**		**Decision Making Index**	
	Full Sample	Complier Subsample	Full Sample	Complier Subsample	Full Sample	Complier Subsample	Full Sample	Complier Subsample	Full Sample	Complier Subsample
	[1]	[2]	[3]	[4]	[5]	[6]	[7]	[8]	[9]	[10]
Age at Marriage	0.233***	0.242***	-0.098***	-0.100***	0.005	-0.002	0.009	0.007	0.003	0.005
	(0.058)	(0.051)	(0.028)	(0.023)	(0.028)	(0.024)	(0.009)	(0.007)	(0.051)	(0.043)
R^2	0.567	0.569	0.415	0.408	0.241	0.25	0.556	0.548	0.176	0.186
First Stage F-statistic	130.00	267.65	128.39	264.52	122.22	240.22	128.47	264.16	116.30	245.22
	[p=0.000]	[p=0.000]	[p=0.000]	[p=0.000]	[p=0.000]	[p=0.000]	[p=0.000]	[p=0.000]	[p=0.000]	[p=0.000]
Kleibergen Paap rK LM statistic	65.07	91.42	64.62	91.1	60.61	81.98	64.59	90.91	62.01	89.57
	[p=0.000]	[p=0.000]	[p=0.000]	[p=0.000]	[p=0.000]	[p=0.000]	[p=0.000]	[p=0.000]	[p=0.000]	[p=0.000]
Observations	10511	9362	10474	9328	9,065	8,127	10,503	9,354	9,309	8,294

Notes: Estimation via TSLS. See main text for description of the outcome variables. All regressions control for women's age, caste affiliation, height, father's years of schooling, mother's years of schooling, number of siblings, and district fixed effects. Standard errors reported in the parentheses are clustered at the district level. ***p < 0.01, **p <0.05, *p < 0.1.

Appendix

Age at Marriage and Women's Labor Market Outcomes in India

Appendix A

A.1 Addressing biases due to selection into labor force

To this point, for the analysis of the effect of women's age at marriage on their hourly earnings, annual wage earnings and work days per year, we have considered only those women who are working. We have excluded nonworking women since we do not observe wages and labor supply for these nonworkers. Given that non-working individuals systematically differ from working individuals, our analysis could be biased. In fact, in accordance to Becker's (1973) theory of specialization, if women drop out of the labor market, ignoring this could potentially lead to underestimation of the beneficial effects of delayed marriage for women. In principle, thus, this might be one reason why we observe statistically insignificant causal effects of delayed marriage on labor market outcomes of women for the full sample and the complier sample. In what follows, we examine whether the observed results and patterns are robust to addressing the selection. Our approach closely follows that used by Wang and Wang (2017).

A.1.1 Selection models and validity of exclusion restriction

To address the selection issue, consider the extended system of Equations (2) and (3) in the presence of endogeneity.

$$MarriageAge_i = \lambda + \delta MenarcheAge_i + \kappa X_i + \eta_i \quad \text{(A1)}$$

$$y_i = \alpha + \beta MarriageAge_i + \gamma X_i + \varepsilon_i \quad \text{(A2)}$$

$$S_i = \mathbb{I}(\delta Z_i - \mu_i \geq 0) \quad \text{(A3)}$$

where S is an indicator and equal to one if one participates in the labor market and zero otherwise; and Z is a vector of exogenous characteristics, which can include a variable not in the set of X. Equation (A3) indicates that a woman or a man decides to participate in the labor market when $\delta Z - \mu \geq 0$. The model can be identified under typical assumptions for IV and selection models and estimated via a variant of the conventional Heckman model (see Wooldridge (2010, p. 809) for details).

As is noted by Wang and Wang (2017) and many others, Heckman type of selection models do not perform quite well even though identification can be achieved through distributional assumption without an exclusion restriction (i.e., $Z = X$). To address this concern, we include an exclusion restriction—spousal earnings (Z)—to aid identification. Specifically, Z equals one if spousal earnings is greater than median income and zero otherwise. This choice of exclusion restrictions for the labor supply equation (particularly for that by females) is a popular one in the literature, and similar variables have been used in the previous literature (e.g. Buchinsky 2001; Chang 2011; Martins 2001). Below, we present our evidence supporting this choice.

To assess the validity of our exclusion restriction, we present two sets of results in Table B5 based on the full sample as well as the complier sample. The first set is concerned with the strength of empirical relationship between our exclusion restriction and labor force participation decision of the women. The literature has generally found strong evidence that spousal income influences a woman's decision to participate in the labor market (e.g. Mroz, 1987; Zabel, 1993). We present the marginal effects of spousal income on a woman's probabilities of labor force participation. Consistent with the literature, our first-stage results show that spousal income indeed has a negative and statistically significant effect on labor force participation rates among women. Specifically, having a spouse who earns more than median income can reduce female labor force participation rate by 12.2% for the full sample (Column 1) by 11.8% for the complier sample (Column 3).

The second set of results is concerned with the independence of an exclusion restriction; the exclusion restriction must be independent of potential labor market outcomes (or conditional on X). Such assumption may be violated if spousal earnings has any direct effects on individual wages or labor supply, or is indirectly related with individuals wages or labor supply through other channels. As noted by Wang and Wang (2017), one possibility is selection into marriage based on unobservable determinants of individual labor market outcomes, which implies potential non-zero correlation between spousal income and the error term as well. To formally test whether this assumption (along with the monotonicity assumption) is violated, we conduct a formal test based on a novel method proposed in Huber and Mellace (2014). They show that under our model assumptions, the following inequalities hold:

$$\begin{aligned} \mathbb{E}[y|z &= 1, S = 1, y_i \leq y_q] \leq \mathbb{E}[y|z = 0, S = 1] \\ &\leq \mathbb{E}[y|z = 1, S = 1, y_i \geq y_{1-q}] \end{aligned}$$

where y_q the qth conditional quantile in the conditional outcome distribution given $Z = 1$ and $S = 1$. Such inequalities imply the following null hypothesis:

$$\begin{aligned} \mathbb{E}[y|z &= 1, S = 1, y_i \leq y_q] - \mathbb{E}[y|z = 0, S = 1] \leq 0 \\ \mathbb{E}[y|z &= 0, S = 1] - \mathbb{E}[y|z = 1, S = 1, y_i \geq y_{1-q}] \leq 0 \end{aligned}$$

Huber and Mellace (2014) propose a test procedure to verify these inequalities. A negative test statistic with a large p value indicates that the IV is valid. The results for the full sample and complier sample are presented in columns (2) and (4) of Table B5 respectively. For both the samples, we fail to reject the validity of our exclusion restriction, strongly in favor of the use of the presence of spousal income as an exclusion restriction for the selection equation. These results, while not necessarily definitive, do increase our confidence in the identification assumption used in our analysis.

A.1.2 Results addressing selection

We now turn to actual estimates addressing the selection issue. We repeat all of our analysis addressing the selection issue. The results for the full sample and complier sample are presented in Table B6. As we can see, all of our baseline results continue to hold. Not only do we find similar patterns in our estimates; we generally find estimates to be remarkably similar in magnitudes as well. Specifically, we again find a statistically insignificant effect of women's delayed marriage on hourly earnings, annual wage earnings, and work days per year for the full sample and the complier sample. Since the IV estimates continue to be statistically insignificant even after correcting for selection bias, we conclude that there does not exist a causal relationship between women's age at marriage and their hourly earnings, annual wage earnings, and work days per year.

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Appendix B

Table B1. OLS estimates of the effect of age at marriage on labor market outcomes, Alternative Complier Subsample

	[1]	[2]	[3]	[4]
	Participation	Hourly Earnings	Annual Wage Earnings	Work Days Per Year
Age at Marriage	-0.010***	0.016***	0.028***	1.676***
	(0.001)	(0.003)	(0.005)	(0.466)
R^2	0.283	0.329	0.420	0.276
Observations	32584	8861	8861	8861

Notes: Estimation via OLS. All regressions control for women's age, caste affiliation, height, father's years of schooling, mother's years of schooling, number of siblings, and district fixed effects. Standard errors reported in the parentheses are clustered at the district level. ***p < 0.01, **p <0.05, *p < 0.1.

Table B2. IV estimates of the effect of age at marriage on labor market outcomes, Alternative complier subsample

	[1]	[2]	[3]	[4]
	Participation	Hourly Earnings	Annual Wage Earnings	Work Days Per Year
Age at Marriage	-0.003	0.005	0.021	0.157
	(0.004)	(0.008)	(0.016)	(1.395)
R^2	0.282	0.327	0.419	0.275
First stage F statistic	1115.42	417.13	417.13	417.13
	[p=0.000]	[p=0.000]	[p=0.000]	[p=0.000]
Kleibergen Paap rK LM statistic	176.25	102.11	102.11	102.11
	[p=0.000]	[p=0.000]	[p=0.000]	[p=0.000]
Observations	32584	8861	8861	8861

Notes: Estimation via TSLS. All regressions control for women's age, caste affiliation, height, father's years of schooling, mother's years of schooling, number of siblings, and district fixed effects. Standard errors reported in the parentheses are clustered at the district level. ***p < 0.01, **p <0.05, *p < 0.1.

Table B3. OLS and IV estimates of the effect of age at marriage on labor market participation, Alternative definitions

	Hours worked in last one year ≥ 240 hours		**Ever participated**	
	OLS	IV	OLS	IV
Age at Marriage	-0.006***	-0.005	-0.002**	-0.006
	(0.001)	(0.006)	(0.001)	(0.007)
R^2	0.252	0.252	0.213	0.213
First Stage F-statistic		300.53		300.16
		[p=0.000]		[p=0.000]
Kleibergen Paap rK LM statistic		124.13		124.01
		[p=0.000]		[p=0.000]
Observations	37,655	37,655	37,609	37,609

Notes: All regressions control for women's age, caste affiliation, height, father's years of schooling, mother's years of schooling, number of siblings, and district fixed effects. Standard errors reported in the parentheses are clustered at the district level. ***p < 0.01, **p <0.05, *p < 0.1.

Table B4. IV estimates of the effect of age at marriage on labor market outcomes: Region wise analysis

	Participation in last one year			
	[1]	[2]	[3]	[4]
	North	Central	South	North-East
Age at Marriage	0.014	0.005	-0.017***	0.016
	(0.015)	(0.011)	(0.006)	(0.038)
R^2	0.260	0.313	0.241	0.102
First Stage F-statistic	60.67	153.40	175.00	15.16
	[p=0.000]	[p=0.000]	[p=0.000]	[p=0.002]
Kleibergen Paap rK LM statistic	33.65	56.95	38.62	7.69
	[p=0.000]	[p=0.000]	[p=0.000]	[p=0.006]
Observations	12,667	15,511	8,022	1,455
	Ever Participated			
	[1]	[2]	[3]	[4]
	North	Central	South	North-East
Age at Marriage	0.025	-0.021	-0.009	0.003
	(0.016)	(0.014)	(0.007)	(0.030)
R^2	0.124	0.246	0.185	0.143
First Stage F-statistic	60.61	152.26	175.25	15.08
	[p=0.000]	[p=0.000]	[p=0.000]	[p=0.002]
Kleibergen Paap rK LM statistic	33.63	56.64	38.66	7.6
	[p=0.000]	[p=0.000]	[p=0.000]	[p=0.006]
Observations	12,666	15,491	8,003	1,449
	Hourly Earnings			
	[1]	[2]	[3]	[4]
	North	Central	South	North-East
Age at Marriage	0.021	-0.014	0.008	0.351
	(0.045)	(0.025)	(0.009)	(0.338)
R^2	0.35	0.268	0.284	-0.722
First Stage F-statistic	19.77	43.17	119.51	0.85
	[p=0.000]	[p=0.000]	[p=0.000]	[p=0.371]
Kleibergen Paap rK LM statistic	15.00	29.99	30.22	0.79
	[p=0.000]	[p=0.000]	[p=0.000]	[p=0.373]
Observations	2,522	4,537	3,205	247

Table B4. IV estimates of the effect of age at marriage on labor market outcomes: Region wise analysis (Continued)

	Annual Wage Earnings			
	[1]	[2]	[3]	[4]
	North	Central	South	North-East
Age at Marriage	0.107	0.002	0.005	0.389
	(0.074)	(0.045)	(0.019)	(0.430)
R^2	0.426	0.426	0.273	0.042
First Stage F-statistic	19.77	43.17	119.51	0.85
	[p=0.000]	[p=0.000]	[p=0.000]	[p=0.371]
Kleibergen Paap rK LM statistic	15.00	29.99	30.22	0.79
	[p=0.000]	[p=0.000]	[p=0.000]	[p=0.373]
Observations	2,522	4,537	3,205	247
	Work Days Per Year			
	[1]	[2]	[3]	[4]
	North	Central	South	North-East
Age at Marriage	1.964	-4.73	0.771	57.695
	(6.941)	(4.129)	(1.538)	(70.792)
R^2	0.346	0.241	0.17	-1.488
First Stage F-statistic	19.77	43.17	119.51	0.85
	[p=0.000]	[p=0.000]	[p=0.000]	[p=0.371]
Kleibergen Paap rK LM statistic	15.00	29.99	30.22	0.79
	[p=0.000]	[p=0.000]	[p=0.000]	[p=0.373]
Observations	2,522	4,537	3,205	247

Notes: Estimation via OLS. North region includes Chandigarh, Delhi, Haryana, Himachal Pradesh, Jammu and Kashmir, Punjab, Rajasthan, Uttar Pradesh and Uttarakhand. Central Region includes Bihar, Chhattisgarh, Dadra and Nagar Haveli, Daman and Diu, Goa, Gujarat, Jharkhand, Madhya Pradesh, Maharashtra, Orissa and West Bengal. South region includes Andhra Pradesh, Karnataka, Kerala, Pondicherry and Tamil Nadu. North-East region includes Arunachal Pradesh, Assam, Manipur, Meghalaya, Mizoram, Nagaland, Sikkim and Tripura. All regressions control for women's age, caste affiliation, height, father's years of schooling, mother's years of schooling, number of siblings, and district fixed effects. Standard errors reported in the parentheses are clustered at the district level. ***p < 0.01, **p <0.05, *p < 0.1.

Table B5. Validity tests of instruments in selection models

	Full Sample		Compliers	
	[1]	[2]	[3]	[4]
	Marginal effects	Validity test	Marginal effects	Validity test
Husband's income	-0.122***	-3.547	-0.118***	-3.587
	(0.010)	[p=1.000]	(0.010)	[p=1.000]

Notes: All regressions control for women's age, caste affiliation, height, father's years of schooling, mother's years of schooling, number of siblings, and district fixed effects. Standard errors reported in the parentheses are clustered at the district level. ***p < 0.01, **p <0.05, *p < 0.1. The validity test of the IV is developed in Huber and Mellace (2011). The null hypothesis is that the IV is valid.

Table B6. Selection-bias corrected IV estimates of the effect of age at marriage on labor market outcomes

	Full Sample			Compliers		
	[1]	[2]	[3]	[4]	[5]	[6]
	Hourly Earnings	Annual Wage Earnings	Work Days Per Year	Hourly Earnings	Annual Wage Earnings	Work Days Per Year
Age at Marriage	0.005	0.014	-0.357	0.005	0.021	0.096
	(0.010)	(0.020)	(1.669)	(0.009)	(0.017)	(1.419)
R^2	0.351	0.441	0.282	0.362	0.444	0.289
First Stage F-statistic	130.05	130.05	130.05	268.41	268.41	268.41
	[p=0.000]	[p=0.000]	[p=0.000]	[p=0.000]	[p=0.000]	[p=0.000]
Kleibergen Paap rK LM statistic	65.07	65.07	65.07	91.41	91.41	91.41
	[p=0.000]	[p=0.000]	[p=0.000]	[p=0.000]	[p=0.000]	[p=0.000]
Observations	10511	10511	10511	9362	9362	9362

Notes: Estimation via TSLS. All regressions control for women's age, caste affiliation, height, father's years of schooling, mother's years of schooling, number of siblings, and district fixed effects. Standard errors reported in the parentheses are clustered at the district level. ***p < 0.01, **p <0.05, *p < 0.1.