Columbia University

Department of Economics
Discussion Paper Series

Female Labor Market Conditions and Family Formation

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Ayako Kondo is a PhD student in the Department of Economics, Columbia University. Faculty members Janet Currie and Till von Wachter have recommended the inclusion of this paper in the Discussion Paper Series.

Discussion Paper No.: 0809-08

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November 2008

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November 7, 2008

Abstract

Slack labor market conditions for women relative to men increase marriage rates for young women. One concern is that this increase may be from marginal marriages due to some females lowering their reservation match quality, and so lead to future divorces and possibly to increases in female headship and poverty. This paper examines the long-term consequences of such marriages using data from the Survey of Income and Program Participation and the Panel Study of Income Dynamics. I find that the marriages induced by relatively poor economic conditions for women reflect shifts in the timing of marriage among young women who would eventually marry anyway. Labor market conditions at age 18-20 do not affect the fraction of women who will marry by age 30. Further, labor market conditions at marriage are uncorrelated with the probability of divorce or with spouses' characteristics, and marrying young in response to labor market shocks does not significantly affect a woman's fertility or labor supply. These findings are consistent with a model in which economic conditions affect women's search intensity without affecting their reservation match quality.

*I am particularly indebted to my advisors, Janet Currie, Till von Wachter and Bentley MacLeod, for their guidance and support. I also wish to thank Shubha Chakravarty, Pierre-André Chiappori, Lena Edlund, Joshua Goodman, Mariesa Herrmann, Fung-Mey Huang, Tümer Kapan, Dongshu Ou, Johannes Schmieder, Herdis Steingrimsdottir, Matthew Wai-Poi, and Jane Waldfogel for their helpful suggestions. I benefited from comments by the participants at student colloquia in Columbia, the Econometric Society European and North American Summer Meetings, the Western Economic Association, the Japanese Economic Association, and the Institute for Statistical Research.

1 Introduction

Most developed countries have experienced a substantial long-run decline in the marriage rate associated with women's improved position in the labor market. This effect also exists in the short run: Better labor market opportunities for women relative to men decrease the fraction of married women in the local labor market (Schultz 1994, Blau, Kahn and Waldfogel 2000). These effects on marriage incidence are often interpreted as changes in the number of women who will ever marry. In particular, many studies on the effect of labor market conditions on marriage formation, including Blau et al (2000) and Schultz (1994), have referred to the increased number of single mothers as a motivation for studying marriage incidence. Yet, whether a decline in marriage incidence for young women really increases the number of single mothers depends on whether it decreases the number of women who will enter stable marriages during their lifetimes. Although the contemporaneous effects of labor market conditions on marriage incidence are well-documented, there remains much to learn about how labor market conditions in youth affect the transitions in marital status in the long run.

This paper examines the long-term consequences of marriages induced by temporary labor market fluctuations. If a negative shock to the female labor market induces some couples who would not marry otherwise to marry, it may increase the probability of future divorce. On the other hand, it is possible that economic conditions only shift the timing of marriage without affecting future divorce rates or other observable characteristics of marriages. Acceleration in the formation of stable marriages has quite different implications for demographic trends and public welfare than an increase in poorly matched, short-lived marriages. Since previous analyses relied on cross-sectional data, they were not able to distinguish between the two. To overcome this limitation, this paper exploits individual women's marriage histories taken from the Survey of Income and Program Participation (SIPP) and the Panel Study of Income Dynamics (PSID), which are linked with unemployment rates by gender at the state-year level.

Using these data, I confirm that young women are more likely to marry when labor market conditions for women are bad relative to those of men. Despite this contemporaneous effect on the marriage rate, gender-specific unemployment rates at the time of marriage are not systematically correlated with the probability of divorce later in life. Also, unemployment rates by gender at age 18-20, when these women enter the marriage market, do not affect the fraction who will marry by age 30 in the affected cohort. These findings suggest that increases in the incidence of marriage due to gender-specific labor market shocks are primarily due to acceleration in the timing of marriage among people who would eventually marry without these shocks and that they are due to increased search effort rather than decreased reservation match quality. I also find that the contemporaneous effects on marriage incidence for women older than 24 are opposite in sign to those for younger women, and that this heterogeneity is not attributable to differences in factors such as educational backgrounds, employment status, and the likelihood of having children before marriage.

Furthermore, I find little evidence that the fertility and the labor supply of women are affected by marrying younger due to gender-specific labor market conditions favorable for marriage. Unemployment rates by gender at marriage are not correlated with the likelihood of having a child or the number of children ever born to a woman. Neither are unemployment rates at age 18-20, despite their significant effects on marriage timing. Moreover, an increase in marriage incidence for young women due to worse female labor market conditions is not accompanied by any changes in the prevalence of single parenthood. In addition, gender-specific unemployment rates in youth do not have significant effects on labor supply and income in the mid-thirties, except that a higher female unemployment rate slightly increases weeks worked.

These findings can be interpreted using a marital search model with endogenous search intensity. Since the probability of marriage is a product of the probability that a woman meets a man and the probability that they agree to marry conditional on meeting, an increase in the incidence of marriage must come either through an increase in search effort or a reduction in reservation match quality which is compensated for by a temporary rise in the relative earnings of men. The negligible effects on the divorce probability imply that women do not lower their reservation match quality in order to increase the probability of marriage. On the other hand, an increase in search effort accelerates the timing of marriage without affecting the likelihood of future divorce. Moreover, if search frictions are modest so that a temporary increase in search effort does not change the probability of meeting a suitable mate over one's lifetime, the fraction of women who eventually marry does not change either.

The rest of the paper is organized as follows. The next section reviews the background literature. Section 3 describes the data, and Section 4 presents the empirical results. Section 5 discusses how and to what extent the empirical findings can be explained by a simple model of marital search. Section 6 gives concluding remarks.

2 Background

Becker (1973) provides a theoretical framework in which, under certain conditions, better labor market conditions for women increase the opportunity cost of marriage and better labor market conditions for men increase the gain from marriage. In his model, a decline in the male wage rate has a negative effect on marriage incidence if the marginal utility from income is higher for a married couple than for a single man[1] and marriage does not reduce men's labor supply. Then, for a decline in the female wage rate to have a positive effect on marriage incidence, a married woman's labor supply must be sufficiently lower than a single woman's to compensate for the higher marginal utility from income for a married couple.[2]

Although this condition does not hold for all women, empirical evidence suggests that this condition holds for the majority of young couples in the United States. Keeley (1975) empirically shows that the correlation between wages and the age at first marriage is positive for women and negative for men. Furthermore, Van Der Klaauw (1996) finds that gains from marriage for a woman are decreasing in her wage rate and increasing in her husband's earnings, using a structural model that addresses the interdependence between marital status and labor supply. Moreover, even though the gender wage gap has narrowed considerably since the 1970s, it is still largely the case that wives spend more time on child-rearing and, on average, the labor supply of married women remains considerably lower than that of unmarried women with the same age and educational background.[3]

[1] This assumption is plausible if a spouse's consumption has a positive externality.

[2] Note that the effects of male and female wage rates are not necessarily symmetric in the magnitude.

[3] Among non-Hispanic white women aged 17-35 in the SIPP 1996 and 2001 panels, married women are 13.8 percentage point less likely to be fulltime employed and have worked for 5.8% less weeks during the reference periods than single women with the same age and educational background living in the same state.

Hence, the gain from marriage increases in employment opportunities for women and decreases in the opportunities for men, as long as marriage reduces female labor supply sufficiently and each spouse's wage rate changes with the labor demand for the corresponding gender. In light of this theoretical prediction, empirical studies have found that better opportunities for women in the local labor market decrease the fraction of ever married young women (Freiden 1974, Preston and Richards 1975, White 1981, Schultz 1994, Blau et al. 2000) while better opportunities for men increase it (Freiden 1974, Schultz 1994, Wood 1995, Blau et al. 2000).[4]

While empirical studies on the effects of gender-specific labor market conditions on marriage incidence date back to the 1970s, Blau et al. (2000) is the most comprehensive one to my knowledge. Using MSA-level panel data constructed from Census 1970-1990, they find that gender-specific labor market opportunities[5] affect the fraction of married young women even after controlling for year- and MSA- fixed effects. This paper departs from the literature by investigating not only the contemporaneous effect of gender-specific labor market conditions on the probability of marriage but also the long-term outcomes of marriages induced by temporary labor market fluctuations. The use of panel data also enables me to distinguish the effects through marriage formation from the effects through marriage dissolution, which studies using the Census inevitably confound. In addition, the annual fluctuations in unemployment rates are less correlated with permanent income than decennial changes taken from the Census. The empirical design is similar to that of Dehejia and Lleras-Muney (2004),[6] in the sense that I examine the relationship between unemployment rates at the time of marriage and long-term outcomes, including selection into marriage.

Although not linked with the effects of labor market conditions on marriage formation, a number of studies have explored the relationship between the timing of marriage and its long-term implications. The pioneering work by Becker, Landes and Michael (1977) establishes the

[4] Also, Loughran (2002) and Gould and Paserman (2003) find that the expansion of male wage inequality in the local labor market decreases the marriage rate.

[5] As an index for labor demand, Blau et al (2000) use weighted average of the relative employment of each industry-occupation group in each area by the national importance of each gender-education-race group in each industry-occupation group. I choose to use unemployment rates rather than the same index as Blau et al (2000) because unemployment rates are easier to understand and less noisy when calculated from the CPS, whose sample size is much smaller than the Census.

[6] Dehejia and Lleras-Muney (2004) investigate the relationship between the unemployment rate at the time of a baby's conception and parental characteristics, parental behaviors, and babies' health.

negative correlation between the age at marriage and the probability of divorce.[7] Yet, it is less clear whether this correlation represents a causal effect. If there is a causal effect, couples who are induced to marry younger by worse female labor market conditions should be more likely to end in divorce. In this sense, this paper is related to the literature on the link between marriage timing and the stability of marriage.

This paper is also related to the literature on the motherhood wage penalty. Waldfogel (1997) has established evidence for substantial wage gap between mothers and non-mothers even after controlling for individual fixed effects and the decreased labor supply due to child rearing. Several studies find that the motherhood wage penalty is larger for women who bear children earlier (Blackburn, Bloom and Neumark 1990, Taniguchi 1999, Miller 2007).[8] Thus, if the shift in the timing of marriage leads to a shift in the timing of first births, marrying young in response to labor market shocks may also affect women's income and labor supply in the long-run.

This paper focuses on legal marriage. Although the practical reason for this is the limited availability of information on premarital cohabitation, there are reasons to believe that legal marriage still matters despite the increase in premarital cohabitation and out-of-wedlock childbearing. First, at least for the majority of non-Hispanic white Americans, cohabitation is a step toward marriage rather than a substitute for marriage. Studies based on National Survey of Families and Households 1987-88 (Bumpus, Sweet and Cherlin 1991, Cherlin 1992, Manning and Smock 1995) show that most cohabitors expect to marry their current partners, and that half of the cohabiting never-married non-Hispanic white couples actually marry within two years. Stevenson and Wolfers (2007) confirm similar expectation towards marriage among cohabiting couples based on data collected in the 2000s. Second, even though it is possible for a couple to raise their children together without being legally married, marriage is the most effective way for a man to commit to the fatherhood of his partner's child (Edlund, 2006). Although this commitment could be re-

[7] One of the most recent follow-up studies is Lehrer (2008), who shows the relationship between age at marriage and marital instability is strongly negative up to the late twenties.

[8] Since the timing of the first birth is endogenous, it may be correlated with the inherent earning capacity. Given this concern, Miller (2007) uses biological fertility shocks as an instrument for the age at first births and finds that motherhood delay leads to a substantial increase in earnings. This may appear to contradict with Hotz, McElroy and Sanders (2005), who show teenage childbearing has few long-term negative effects using miscarriage as an instrument. Yet, this apparent contradiction is likely to be attributable to heterogeneity in the effects across different population, since the motherhood wage penalty is larger for high skilled women (Ellwood, Wilde and Batchelder, 2004).

pealed by divorce, divorce is much more costly than the separation of unmarried couples. Finally, more than 80 percent of non-Hispanic white American women marry by the age of 30, and more than 80 percent of first births by non-Hispanic white mothers occur after first marriage. Albeit on the decline, legal marriage is still an important step in the family formation of American women.

3 Data

3.1 Individual Level Data: SIPP and PSID

The main source of individual women's data is the Survey of Income and Program Participation (SIPP) 1990-2004 Panels.[9] The SIPP is a series of short panel surveys conducted by the Census Bureau, with sample sizes ranging from approximately 14,000 to 36,700 households. Although each SIPP core panel covers at most four years, supplementary topical modules provide rich retrospective information. I use the Marriage History and Migration History Topical Modules attached to the Wave 2 of each panel to construct panel data of the marital status and the state of residence going back to the 1970s. I pool seven panels from 1990 to 2004 and use the appropriate sample weights to address the different sample design between pre- and post- 1996 panels.[10]

The SIPP's sample size is larger than most of the other datasets that have information on marriage histories. Another important advantage of the SIPP is that sample attrition due to divorce is virtually non-existent, because the survey does not actually trace people every year.[11] At the same time, the available retrospective information is limited: employment status and income are available only for the period covered by the core panel, and information on spouses

[9] I restrict my sample to women because some information including fertility history is available only for women. Since the SIPP is a household survey, most of the married men in the data are spouses of women in the same data. The results on marriage formation and dissolution (section 4.1) are confirmed to be the same when estimated with the sample of men.

[10] While households in panels prior to 1996 are taken from the representative US population, from 1996 onward households living in high poverty areas are oversampled. Another important difference is that panels prior to 1996 overlap each other. Until 1993, A new panel had been introduced every year. Then, the 1996 redesign replaced it with larger, non-overlapping panels. Nevertheless, results from 90-93 panels are similar to those estimated with 1996-2004 panels. Therefore, I believe that the change in sampling scheme does not bias the results in this paper.

[11] Strictly speaking, divorced women are about 3-4 percentage points more likely to lack information on the state of residence at the time of marriage than those who are still in their first marriage. They are dropped not only from regressions on gender-specific unemployment rates at marriage, but also from the marriage hazard regression on the contemporaneous unemployment rates. I believe this is less problematic than attrition after marriage, which is typical in longitudinal household data such as the Panel Study of Income Dynamics, because I can include all marriages that contribute to the estimated effects on the marriage hazard in the subsequent analyses.

is also available only if the marriage is still intact at the survey date. Also, since the Fertility History Topical Module asks the year of birth of only the first and last children, it cannot be determined whether a woman's second child was born before she was at a certain age if she has three or more children at the survey date. Another issue is that the state of residence in a given year cannot be determined if the respondent has moved across states twice or more since that year; the appendix provides more detail on this issue.

I also use the Panel Study of Income Dynamics (PSID) for supplemental analyses. The PSID started in 1968 as a nationally representative sample of households in the United States and has tried to trace all households formed by members of the original sample households, as well as all new individuals who joined or were born to the sample households. The PSID contains richer information on the spouse's characteristics than the SIPP, although the PSID's sample size is smaller.

I focus on the first marriage of non-Hispanic white women in the contiguous United States to avoid the issue of selectivity into second marriages and complications arising from differences in social norms about marriage across ethnic groups. Also, I drop women who had married by 1978 because unemployment rates by gender can be calculated at the state level only after 1977. Further, since the majority of women marry by their early twenties, the sample is restricted to women born during the period 1956-1980.

Table 1 presents summary statistics for women born in 1956-1980. Column A is the SIPP sample weighted by the sampling weights and Column B is the PSID sample. The SIPP sample is slightly older and more likely to have "some college" education, but the age at first marriage and the age at first births are reasonably similar. Figure 1 shows the transitions in marital status of women in the SIPP sample. More than 80 percent of the women marry at some point between age 17 and 35, and about half have married by 23. Since very few women marry at age 16 or younger, I drop them from the sample. Also, since fewer cohorts are followed up to older ages, marriages after age 35 are omitted and those who had not married until 35 are treated as if never married.

3.2 State-Year Level Unemployment Rates by Gender

The individual-level data from the SIPP and the PSID are augmented by unemployment rates by gender at the year-state level,[12] calculated from the monthly basic Current Population Surveys. The universe is the non-Hispanic white civilian labor force aged 15-40. I take the annual average to reduce sampling errors. Unemployment rates by gender are assigned to each woman based on her state of residence in each year. Also, to avoid problems arising from the multicollinearity between male and female unemployment rates (see Section 4 for detail), I use the gap between male and female unemployment rates instead of the male unemployment rate.

I use unemployment rates by gender at the state level as proxies for gender-specific labor demand in the local labor market. One of the main sources of differential shocks to labor demand for each gender is changes in industry- and occupation- specific labor demand. It is true that the unemployment rate picks up changes in labor supply as well as labor demand. Nevertheless, the bias on the estimated effects of the unemployment rates on marriage formation is expected to be small. Although marriage could change the labor force status of some women, its impact on the aggregate unemployment rate would be quantitatively negligible because only about 4 percent of women aged 15-40 marry each year.[13] Also, since marriage induces both transitions from employment to unemployment and from in the labor force to out of the labor force, whether marriage raises the female unemployment rate is not clear in theory. Moreover, the actual unemployment rate of married women tends to be similar to that of unmarried women of the same age categories, at least for 15-40 year old non-Hispanic white women in the CPS 1978-2003.

It also is true that, according to Becker's (1973) theory, what really matters is the wage rate the woman could earn if employed rather than the unemployment rate. However, it is practically difficult to construct a gender-specific wage index at the state level because of substantial non-employment among married women. I could construct an index following Blau et al (2000), which is a weighted average of employment share of industry-occupation in each state by share of gender

[12] Maine and Vermont are grouped together, and North Dakota, South Dakota and Wyoming are grouped together, because the original variable for the state of current residence in the SIPP is defined in such a way. Alaska, Washington DC, and Hawaii are dropped.

[13] There are about total 2,200,000 marriages per year (Vital Statistics), 90% of which involve women younger than 45 (Census Bureau's report based on SIPP 2001). The population of 18-39 years old women is about 50,000,000 (Census 2000).

in national employment of each industry-occupation. However, such an index does not seem to have any advantage to the unemployment rate in terms of the relationship with the wage rate. Also the industry-occupation composition is as endogenous to labor supply as the unemployment rate. Moreover, given the smaller sample size of the CPS than the Census, which is used by Blau et al (2000), the unemployment rate is less noisy and intuitively easier to understand. Thus I choose gender-specific unemployment rates at the state level as the best available proxy for annual fluctuations in the gender-specific labor markets.

Table 2 reports summary statistics for unemployment rates by gender at the state-year level. Since identification is based on variations net of state- and year- fixed effects, I report the residuals as well as the raw rates. About half of the variation in the female unemployment rate remains after controlling for state- and year- fixed effects and the male-female gap in the unemployment rates. Variation in the male-female gap is also substantial. Figure 2 plots the female unemployment rate and the male-female gap for the United States and five randomly picked states to confirm that there are substantial variations in trends and cycles of gender unemployment rates across states.

3.3 Descriptive Analysis

Before going into more formal analysis, it is informative to look at the correlation between the female unemployment rates in youth and young women's subsequent marriage/fertility behavior in later years. This analysis is done at the cohort level. A cohort is defined as a group of women who were born in the same state and year. Figure 3 presents scatter plots of various cohort level outcomes over the average female unemployment rate of the state in the years when each cohort was 18-20 years old.[14] Each cohort is weighted by the sum of the sample weights of women in it.

First, panel A shows a positive correlation between the female unemployment rate at age 18-20 and the fraction who have married by age 22. That is, women who experienced worse labor market conditions in their late teens are more likely to marry by their early twenties. This is consistent with the existing evidence that worse female labor market conditions increase the incidence of marriage for young women.

[14] As a robustness check, replacing unemployment rates at age 18-20 with those at age 19-21 do not qualitatively change the results of any analyses in this paper.

In panel B, however, the female unemployment at age 18-20 is uncorrelated with the fraction of who have ever married by the mid-thirties. This suggests that the positive correlation in panel A is primarily driven by shifts in marriage timing for those who would marry by age 35 anyway. Furthermore, as shown in panel C, the female unemployment rate at age 18-20 is also uncorrelated with the probability of having a child by age 35. Women who experienced worse labor market conditions for women in youth are no more likely to have a child even though they tend to marry earlier.

Female labor market conditions in youth are uncorrelated with the fraction who will ever marry or ever have a child in the affected cohort, but what about the number of single mothers? I define a "single mother" as a woman who had a child at least a year prior to marriage or divorced after having a child. I then calculate the fraction of women who have ever experienced this "single mother" status by age 35 for each state. Panel D of Figure 3 plots this variable over the female unemployment rate at age 18-20. Female labor market conditions in youth are not correlated with the likelihood of becoming a single mother, despite the perception that changes in marriage incidence due to labor market fluctuations affect the prevalence of single parenthood.

4 Empirical Model

The empirical analysis begins by confirming of the effects of contemporaneous unemployment rates by gender on the probability of getting married. Then, to see whether marriages induced by gender-specific labor market conditions are poorer matches than those formed without such shocks, I estimate the effects of gender-specific unemployment rates at the time of marriage on the probability of divorce and on observable characteristics of spouses. Next, to see whether the increase in marriage incidence for young women results in an increase in the fraction who eventually marry in the cohort, I examine the effects of gender-specific unemployment rates that a woman experienced at age 18-20 on the likelihood of having ever married at each age; 20, 22, 24, 26, 30, 35. If the change in marriage incidence for young women is a timing effect, gender-specific unemployment rates experienced at age 18-20 will not affect the likelihood of having ever married by their thirties.

I exploit variations in gender-specific unemployment rates across states over year to control for state fixed effects and nation-wide year effects. Standard errors are estimated by clustering by states to take into account autocorrelated state-specific random shocks. Also, since the male unemployment rate is positively correlated with the female unemployment rate and their coefficients are expected to be opposite in sign, including both male and female unemployment rates as explanatory variables would boost the absolute value of both coefficients. Therefore, I use the gap between male and female unemployment rates instead of the male unemployment rate. The coefficient of the female unemployment rate is interpreted as the overall effect of local labor market conditions which the woman faces, and that of the male-female gap should be the effect of labor market conditions for her prospective spouse relative to those for herself.

To estimate the contemporaneous effects of gender-specific unemployment rates on marriage incidence, I specify the following Cox's proportional Hazard model:

$$M_{its} = \lambda(age_{it}; birthy) \exp(\alpha_{age} u^w_{ts} + \beta_{age}(u^m_{ts} - u^w_{ts}) + \eta_s + \varepsilon_{its}) \tag{1}$$

where M_{its} is the probability of getting married during the calendar year t for a woman i living in state s, age_{it} is woman i's age in year t, and u^w_{ts} and u^m_{ts} are female and male unemployment rates in year t in state s. The baseline hazard λ depends on women's age and is stratified by the year of birth (*birthy*). This is equivalent to controlling for calendar-year*birth-year fixed effects. η_s is a state fixed effect, and ε_{its} is the remaining error.

Since marriage incidence varies with women's age and the value of marriage is likely to change with women's age, the effects of gender-specific unemployment rates on marriage incidence may well vary with women's age. Therefore, I allow α and β to depend on women's age by taking interactions with dummies for each single year age or four age categories (17-20, 21-23, 24-27, 28-35). Also, since more educated women tend to have stronger aims for careers, they could respond to labor market fluctuations in a different way than less educated women. Thus, I estimate (1) by educational background subsamples.

To examine whether increased marriage incidence is associated with a higher probability of future divorce, I estimate the following probit model:

$$\Pr(\text{Divorce in 5 or 10 years}) = \Phi(\alpha u^w_{TS} + \beta(u^m_{TS} - u^w_{TS}) + \eta_S + \xi_T + \varepsilon_{iTS}) \quad (2)$$

where T is the year of marriage, S is the state of residence during the year of marriage, u^w and u^m are female and male unemployment rates, η_S is a dummy variable for the state of marriage, and ξ_T is a dummy variable for the year of marriage.

Since the timing of marriage is endogenous, I also estimate the effects of the female unemployment rate observed during a woman's youth (18-20 years) in her state of birth. These gender-specific unemployment rates experienced at age 18-20 are expected to affect the probability of marriage by age 20. Accordingly, dummy variables for the year of marriage and for the state of marriage are replaced with dummy variables for the year of birth and for the state of birth.

In addition, I estimate the following Cox's hazard model with controls for the contemporaneous unemployment rates:

$$D_{itTS} = \lambda(t-T;T)\exp(\alpha u^w_{TS} + \beta(u^m_{TS} - u^w_{TS}) + \gamma u^w_{ts} + \delta(u^m_{ts} - u^w_{ts}) + \eta_S + \varepsilon_{itTS}) \quad (3)$$

where D_{itTS} is the divorce hazard, or the probability of getting divorced, and the baseline hazard depends on years since marriage $(t-T)$ and is stratified by the year of marriage, which is equivalent to controlling for year-of-marriage*calendar-year fixed effects.

Note that α and β in (2) and (3) do not imply causal effects on the likelihood of divorce for otherwise identical couples. The purpose of estimating (2) and (3) is to examine whether gender-specific labor market conditions induce couples who are likely to divorce in future to marry, i.e. the selection into marriage. Yet, it is true that gender-specific labor market conditions at the time of marriage can have a causal effect on the likelihood of divorce. In particular, worse labor market conditions for women are expected to encourage investments in marriage specific capital by lowering the opportunity cost. Hence, the increase in marriage specific investments may offset a lower match quality, causing the divorce probability to remain the same even if women do change their reservation match quality in response to labor market conditions. In light of this prediction, I investigate the relationship between unemployment rates by gender at the time of marriage and long-term outcomes such as the number of children and labor supply later.

As an alternative way to assess the effects on selection into marriage, I estimate the correlations between the observable characteristics of spouses and gender-specific unemployment rates at the time of marriage using linear OLS and probit models with the same set of explanatory variables as in (2). Like in the probit model for the divorce rate, I also estimate the effects of gender-specific unemployment rates at age 18-20, replacing marriage year- and state- fixed effects with birth year- and state- fixed effects.

Furthermore, if some couples who would not marry without such shocks are induced to marry, gender-specific unemployment rates in youth must affect the fraction who will ever marry. Thus, I estimate the effects of unemployment rates by gender experienced at age 18-20 on the future marital status. I begin by the median regression of the age at first marriage as follows:

$$Agemar_{i\tau s} = \alpha \bar{u}^{w}_{\tau s} + \beta(\bar{u}^{m}_{\tau s} - \bar{u}^{w}_{\tau s}) + \eta_s + \xi_\tau + \varepsilon_{i\tau s} \tag{4}$$

where $\bar{u}^{w}_{\tau s}$ and $\bar{u}^{m}_{\tau s}$ are the average female and male unemployment rates in years when the woman was 18-20 years old in the state of birth, η_s is a dummy variable for the state of birth, ξ_τ is a dummy variable for the year of birth, and $\varepsilon_{i\tau s}$ is the remaining error. Equation (4) is estimated with women in SIPP 2001 and 2004 panels born in 1960-1970, so that I can track each woman at least up to age 30. *Agemar* of a woman who has never married is assumed to be above the median.

Next, I estimate the effects on the likelihood of having ever married by a specific age using the cohort-level dataset used in Figure 1. The dataset is restricted to women in SIPP 2001 and 2004 born in 1960-1970 to prevent picking up cohort-level trends, and is collapsed by the year and state of birth. I estimate the following linear model separately for each age using this collapsed dataset:

$$\%married = \alpha \bar{u}^{w}_{\tau s} + \beta(\bar{u}^{m}_{\tau s} - \bar{u}^{w}_{\tau s}) + \eta_s + \xi_\tau + \varepsilon_{i\tau s} \tag{5}$$

Each cohort is weighted with the sum of the sample weights of the women in the year-state cell.

After these analyses of marriage formation, I investigate the effects of gender-specific labor market conditions at marriage and in youth on the fertility and the labor supply in the long run. Specifically, I begin by estimating the effects on the number of children born by age 35 with a linear OLS model, with the same set of explanatory variables as in equation (2). Like in the model for the divorce rate and the husband's characteristics, I also estimate the effects of gender-specific unemployment rates at age 18-20, replacing marriage year- and state- fixed effects with birth year- and state- fixed effects. Next, I examine the effects of gender-specific unemployment rates in youth on the age at birth of the first child using the median regression exactly the same as (4) except for the dependent variable. Then, I estimate the effects of gender-specific unemployment rates in youth on the probability of having a child by age 25, 30, and 35 in the same way as in (5). I also estimate the effects on the probability of entering the single motherhood, which is defined as having a child prior to marriage or divorcing after having a child.

Since marrying young in response to labor market conditions in youth may affect the woman's earnings and labor supply, I estimate reduced form effects of gender specific unemployment rates at age 18-20 on various economic outcomes in the mid-thirties. It is true that the estimated coefficients of the unemployment rates may not be interpreted as causal unless the effects of labor market conditions in youth on women's investment in human capital through other channels can be ruled out.[15] However, it is even harder to estimate the long-term effects of gender-specific unemployment rates at marriage because the timing of marriage is endogenous and correlated with women's preference for labor supply and productivity and the presence of an infant child. Thus, as suggestive evidence, I estimate the following linear regression:

$$Deptvar_{i\tau s} = \alpha \bar{u}^{w}_{\tau s} + \beta(\bar{u}^{m}_{\tau s} - \bar{u}^{w}_{\tau s}) + \gamma educ_i + age_i + \eta_s + \xi_\tau + \varepsilon_{i\tau s} \tag{6}$$

where *Deptvar* represents economic outcomes such as log personal earnings and weeks worked last year, $\bar{u}^{w}_{\tau s}$and $\bar{u}^{m}_{\tau s}$are the average female and male unemployment rates in years when the woman

[15] Although such effects cannot be ruled out, their quantitative impact seems to be negligible. Kondo (2007) shows that the effects of a recession at entry to the labor market on wages and employment are negligible for non-Hispanic white women, in contrast to persistent, significantly negative effects for men. Since Card and Lemieux (2000) shows that the unemployment rate at age 18 has a slightly negative effect on the years of schooling, all regressions shown in Table 9 include controls for educational background.

was 18-20 years old in the state of birth, *educ* is her educational background, age_i is a dummy variable for age by single years, η_s is a state-of-birth fixed effect, ξ_τ is a year-of-birth fixed effect, and $\varepsilon_{i\tau s}$ is the remaining error. Since earnings and employment status are available only from the core panel, I restrict the sample to 33-37 year old women.

5 Results

5.1 Effects of Gender-Specific Unemployment Rates on Marriage Formation

Table 3 reports the estimated α and β in equation (1), the coefficients of the female unemployment rate and the male-female gap in the Cox's proportional hazard model for marriage. Panel A shows the estimated effects without allowing them to vary with age. The first column is estimated with the entire sample of the SIPP, and the second to fourth columns are estimated by educational background subsamples. The point estimate of the coefficient of the female unemployment rate implies that 1 percentage point rise in the female unemployment increases the number of total marriage by 4.9 percent, which is equivalent to 0.4 percentage point increase in the probability of getting married for a single women aged 17-35.[16] Likewise, the effect of 1 percentage point rise in the male-female gap decreases the number of total marriage by 3.5 percent and the probability of getting married by 0.3 percentage point. The positive effect of the female unemployment rate and the negative effect of the male-female gap is consistent with existing studies including Blau et al (2000).

The second to forth columns are estimates by educational background subsamples. The positive effect of the female unemployment rate is stronger for more educated women, and the negative effect is stronger for less educated women. This may be because the employment status of the potential husbands of less educated women is more vulnerable to aggregate labor market conditions. The last column shows the estimates from the PSID. The effects are statistically insignificant and smaller than the estimates from the SIPP probably due to the small sample size. The signs of coefficients are consistent.

To see how the effects change with women's age, Panel B presents the estimated coefficients

[16] The average probability of getting married for 17-35 year old women in the SIPP sample is 8.8%.

of unemployment rates interacted with four age categories. The first column, estimated with the entire sample of SIPP, shows that the positive effect of the female unemployment rate on marriage incidence is stronger for younger women and that the effect diminishes with woman's age and becomes significantly negative for women aged 24 or older.[17] The negative effect of the male-female gap in the unemployment rate also diminishes with women's age. Figure 4 plots the effects of 1 percentage-point changes of the female unemployment rate and the male-female gap on the probability of getting married by single year ages. The effect of the female unemployment rate on the marriage hazard is decreasing in age, and the effect of the female unemployment rate turns negative when the effect of the gender gap becomes zero.

The negative effect of the female unemployment rate on the marriage hazard for older women is not attributable to different effects for different educational categories. The second to the fourth columns of Table 3 show a robust pattern within each educational category. Moreover, the age group where the coefficients flip signs is later for more educated women. The last column is estimated with the PSID and looks similar to the results from the SIPP. Section 5.3 further examines potential factors correlated with women's age and may affect the impact of gender unemployment rates on marriage incidence.

One might suspect that changes in marriage incidence could affect the female unemployment rate. As mentioned in Section 3.2, such reverse effects seem to be quantitatively negligible. Moreover, appendix section B1 shows that the estimated effects of unemployment rates by gender on marriage incidence are qualitatively similar to the results of Table 3 when the state unemployment rates are instrumented with the weighted average of nation-wide, industry-occupation specific unemployment rates. Therefore, I believe that any reverse causality of marriage incidence on contemporaneous unemployment rates is negligible.

Since worse female labor market conditions increase marriage incidence for young women, the next question is whether this increased marriage incidence is associated with a higher probability

[17] The results for older woman may look inconsistent with Blau *et al.* (2000)'s findings that better labor market for men and worse labor market for women increase marriage incidence for 25-34 years old women. One possible reason is that their results for older woman pick up the effects of labor market conditions which they experienced when they were young. For example, the proportion of currently married women in women who were 25-34 years old in 1980 Census could reflect any shocks that affected 20-year-old women in the early 1970s. As shown in Figure 5 and Table 6, the cumulative marriage rate of cohorts who experienced relatively worse labor marked conditions for women in their late teens remain higher until the late twenties.

of future divorce. The first and third columns of panel A of Table 4 report the estimates of equation (2), the marginal effects of the female unemployment rate and the male-female gap in unemployment rates at marriage on the probability of divorce within five years and ten years, respectively. Both coefficients are almost zero and not statistically significant.

The second and fourth columns of Table 4 show the effects of unemployment rates by gender at age 18-20. Gender-specific unemployment rates in the state and year of marriage in (2) are replaced with those at age 18-20 based on the state and year of birth, and accordingly, marriage year- and state- fixed effects are replaced with birth year- and state- fixed effects. Again, coefficients are almost zero and statistically insignificant, except for the positive effect of the male-female gap on the 10-year divorce rate. Given the negative effect of the male-female gap on marriage incidence shown in Table 3, the result does not support the hypothesis that marriages induced by labor market fluctuations are more likely to end in divorce.

Since the effects of gender specific unemployment rates on the marriage hazard vary with women's age, I allow coefficients of gender-specific unemployment rates at marriage to vary with the wife's age at marriage.[18] Panel B shows the result; although a few coefficients are statistically significant, there seems to be no systematic relationship between gender-specific unemployment rates at marriage and the likelihood of divorce in the subsequent five years. At least, there is no evidence that an increase in marriage incidence leads to an increase in future divorces.

Further, panel C reports the estimates of equation (3), the Cox's proportional hazard model for the duration of marriage. The estimated coefficients of gender-specific unemployment rates are almost zero, confirming that gender-specific unemployment rates at marriage are not systematically correlated with the duration of marriage. This also shows that contemporaneous *aggregate* labor market conditions do not affect the probability of divorce much. This does not contradict with the findings by Weiss and Wills (1997) that a negative shock to husband's earnings and a positive shock to wife's earnings increase the risk of divorce, since earning shocks at the individual level can convey information about the partner's non-economic suitability as a mate (Charles and Stephens, 2004).

[18] For Tables 5 and 7, which also show effects of unemployment rates at marriage, Appendix Tables B2 and B3 show the effects by age at marriage.

As discussed in Section 4, the divorce rate may be affected by factors other than the inherent match quality. Thus, Table 5 reports the correlations between the observable characteristics of spouses and gender-specific unemployment rates at the time of marriage, as an alternative way to assess whether marriages induced by labor market shocks are poorer matches.[19] I restrict the sample to women in the SIPP core panel who married within 3 years of the survey and also use the PSID, because the SIPP provides information on spouses only for couples who are still married. Dependent variables are the difference between husband's age and the wife's age, the husband's years of schooling, and for the PSID, whether the husband is of the same religion or race. I use both unemployment rates at actual marriage and age 18-20 in the same way as in Table 4.

Panel A of Table 5 shows that women who marry under worse labor market conditions for women have spouses slightly less educated and less likely to be of the same religion and race, although the estimated coefficients are noisy and often inconsistent between the SIPP and the PSID. Yet, the size of the effect seems to be small; for example, 1 percentage point rise in the female unemployment rate at marriage is associated with 0.06 fewer years of schooling in the SIPP sample. Moreover, as shown in panel B, unemployment rates by gender at age 18-20 have no significant effects on the ages and educational backgrounds of husbands. Since the timing of marriage is endogenous, the slightly negative correlation between the female unemployment rate at marriage and husbands' characteristics may rather suggest the negative selection of women who marry in response to worse female labor market conditions. Overall, even though labor market conditions at marriage are weakly correlated with husbands' characteristics, these correlations do not seem to be strong enough to affect the divorce probability.

So far, it has been shown that relatively worse labor market conditions for women increase marriage incidence among women in their teens and early twenties, and that these extra marriages are not associated with a higher probability of divorce in future. Then, does this increase in marriage for young women affect the fraction who will ever marry in the affected cohort? Figure 3 has already shown that the female unemployment rate at age 18-20 is not correlated with the

[19] In general, couples with similar characteristics are less likely to divorce. Becker (1973) presents a theoretical model in which gains from marriage are larger for couples with similar traits. In particular, similarity in religion is an important factor for marital stability (Becker et al 1977, Lehrer and Chiswick 1993). In addition to religion, Weiss and Wills (1997) find that similarity in ethnicity also reduces the probability of divorce.

fraction who have ever married by age 35. Figure 5 further compares the transitions in marital status of women in cohorts who are supposed to marry younger with the rest of the sample. Specifically, I split women born in 1960-1970 in the SIPP 2001 and 2004 panels[20] into two groups, one of which consists of cohorts (defined by state and year of birth) that experienced a female unemployment rate higher than the median (7.6 percent) when they were 18-20 years old. Then, I calculate the fraction who have ever married by each age for each group and plot it over age. Although women in cohorts that experience a higher female unemployment rate and a lower male-female gap at entry to the marriage market are more likely to have married by their early twenties, these differences fade away by their thirties. Thus, worse labor market conditions for women and relatively better labor market conditions for men accelerate the timing of marriage without affecting the fraction of women who will ever marry.

To confirm that these results are robust to controlling for state- and year- fixed effects, first, I estimate the effects of unemployment rates by gender experienced at age 18-20 on the median age at marriage with the median regression defined by equation (4). Women who have never married are assumed to be above the median. The first column of Table 6 shows that women who experienced worse labor market conditions in youth marry younger. Also, relatively worse labor market conditions for men delay marriage.

The rest of Table 6 reports the estimated effects on the likelihood of having ever married by a specific age with in the cohort-level linear model (5). The positive effect of the female unemployment rate at age 18-20 on the fraction of women who have ever married fades away as the cohort ages. The negative effect of the male-female gap also fades away, although the coefficients are not statistically significant even for the early twenties.

To summarize, a high female unemployment rate and a low male unemployment rate increase the incidence of marriage for women younger than early twenties but decrease it for older women. Younger women adjust the timing of marriage according to labor market conditions, but the probability of divorce in future is not affected. Labor market conditions in the teens and early twenties do not affect the fraction of women who will ever marry by the mid-thirties, either. Since the effects of contemporaneous gender-specific unemployment rates on the divorce hazard are also

[20] I restrict my sample so that I can follow the same people from 18 to 30 years old.

weak, the well-documented effects of male and female labor market conditions on the fraction of women who are currently married can be attributed to changes in the timing of marriage for women who would eventually marry anyway.

5.2 Implications for Fertility, Labor Supply and Income

The response of the marriage rate to changes in gender-specific unemployment rates is a timing effect for those who would eventually marry without such shocks. Even so, the timing of marriage itself may affect women's fertility and labor supply decisions. Also, since worse labor market conditions for women may cause some newly married women to withdraw from the labor force, labor market conditions at the time of marriage may affect investments in marriage specific human capital in the long run. Therefore, this section examines whether gender specific unemployment rates at marriage and in youth have long-term effects on women's fertility, labor supply, and income.

First, Table 7 shows the effects of gender-specific unemployment rates at marriage and in youth on the number of children born to each woman by her mid-thirties. Since the SIPP does not provide the years of birth of all children and asks the number of children only at the Wave 2 interview, the number of children born to a woman by a specific age is available only for those who are at that age at the time of survey. Therefore, the sample used in the first two columns of Table 7 is restricted to women who are 35-37 years old at the Wave 2 interview. The first column shows negligible correlations between unemployment rates by gender at marriage and the number of children. It suggests that marrying in response to gender-specific labor market conditions do not affect the fertility in the long-run.

It is true that the estimated coefficients of gender-specific unemployment rates at marriage may not be causal effects because these coefficients may be affected by the selection into marriage. Yet, the direction of the causal effects and the spurious effects from selection are expected to be the same. On one hand, worse labor market conditions for women are expected to induce some women to withdraw from the labor force. On the other hand, women who want to have children are more likely to be induced to marry because worse labor market conditions for women lower the opportunity cost of bearing a child. Since the estimated coefficients are not statistically distinct from zero, both the causal and selection effects are negligible.

Furthermore, the second column of Table 7 shows insignificant effects of unemployment rates at age 18-20 for all women including those who have not married by age 35. That is, even though worse labor market conditions for women accelerates the timing of marriage, the shift in marriage timing does not significantly affect the number of children that these women will have by their mid thirties. Since the sample restriction imposed on the SIPP sample is complicated, the last column shows the effects of unemployment rates at marriage estimated with the PSID as a robustness check.[21] It confirms that the effects are statistically insignificant.

Since young women marry earlier if they face a high female unemployment rate and a relatively low male unemployment rate, these earlier marriages may lead to earlier entry into motherhood. Thus, I estimate the effects of gender-specific unemployment rates at age 18-20 on the age at birth of the first child. The first column of Table 8 shows that worse female labor market conditions in youth do not accelerate the timing of first births, despite their significant effect on marriage timing, which is shown in Table 6. Instead, the male-female gap significantly delays the timing of the first births. The second to fourth columns show the effects of unemployment rates by gender at age 18-20 on the probability of having a child by a certain age using the same specification as Table 6. Labor market conditions in youth do not affect the probability of eventually having a child either, although relatively worse labor market conditions for men seem to delay the timing.

The last two columns of Table 8 show that worse female labor market conditions do not affect the probability that a woman in the affected cohort becomes a single mother. First, even though worse female labor market conditions accelerate the timing of marriage, they do not significantly reduce the probability of having a child before marriage. The dependent variable of the last column is one if the woman either had a child before her first marriage or divorced after the birth of her first child by age 35. Again, unemployment rates at age 18-20 do not have significant effects. The result seems to be consistent with the negligible effects on the divorce probability shown in Table 4.

Even though it does not affect fertility, marrying young in response to labor market conditions

[21] The number of children is observed at the year when the woman becomes 35 (and 36 for 1997-2003 because the survey has become biennially since 1997). I did not estimate the effects of unemployment rates at age 18-20 because the state of birth was not available for too many women in the sample.

in youth may affect the woman's earnings and labor supply. Thus, I estimate reduced form effects of gender specific unemployment rates at age 18-20, which affect the timing of marriage at the cohort level, on various economic outcomes in the mid-thirties. The correlation between labor market conditions in youth and labor supply can provide suggestive information about the effects of changes in marriage timing due to labor market conditions.

Table 9 shows the estimated coefficients of gender-specific unemployment rates at age 18-20 in the linear model defined by equation (6). Panel A presents the estimates for all women including those who are not married. Except for the positive effect of the female unemployment rate on weeks worked last year, labor market conditions at age 18-20 do not affect women's income and employment significantly. The last column confirms that the result of Table 6 holds for this sample, i.e. gender specific unemployment rates at age 18-20 do not affect the fraction of women who marry by their thirties. Panel B is estimated with married women only, which confirms a similar pattern. Moreover, the last column shows that labor market conditions in youth are not correlated with the husband's income either. Overall, labor market conditions in youth do not have significant effects on personal earnings, household income or spouse's earnings, despite the significant effects on the timing of marriage. If there are any effects, worse labor market conditions for women seem to slightly increase the weeks worked.

Years since marriage is correlated with the presence of an infant child, which affects labor supply and perhaps also wages, and the timing of marriage is endogenous and correlated with women's preference for labor supply and productivity. Therefore, it is hard to estimate the long-term effects of gender-specific unemployment rates at the time of marriage on the woman's labor supply and income. Yet, the reduced form effect presented in Table 9 shows no supportive evidence that worse female labor market conditions at marriage reduce women's labor supply.

5.3 Alternative Explanations for Differential Effects across Age

Table 3 and Figure 4 have shown that the female unemployment rate has a negative effect on the probability of getting married for older woman, in contrast to a positive effect for younger women. This section examines several factors that could produce a negative effect of the female unemployment on the marriage hazard and are correlated with age.

First, the classical argument that men have a comparative advantage in market work may no longer be true for women who have settled in stable employment. Table 3 has already shown that the heterogeneous effects across women's age are not attributable to heterogeneity across educational background, which suggests that the negative effect does not come from stronger *aims* for careers. Still, it is worth investigating whether the negative effect appears among women who have not finished their transition to full-time employment.

Therefore, I restrict the sample to women who graduated from college recently. Given the high turnover rate in the entry level labor market in the United States, these women are likely to be still transitioning to stable employment. The first column of Table 10 replicates the first column of Table 3 as a benchmark. The second column shows the estimated coefficients from exactly the same specification except that the sample is limited to women who had at least some college education and completed that education sometime between the year of observation and 3 years prior to it.[22] It does not look much different from the first column. Using the information about employment status and enrollment status in the core panel, I also estimate the same model with the sample of women who are not employed full-time or enrolled full-time in the previous year.[23] Column (3) of Table 10 shows that the effect of the female unemployment rate on marriage incidence flips signs in the mid-twenties even for women who are not employed full-time or enrolled full-time, although the negative effect are not statistically significant due to the standard errors boosted probably by the small sample size. Therefore, the more stable employment does not seem to be able to explain all of the negative effect of the female unemployment rate for older women.

Second, even unmarried, older women are more likely to have a child. If men do not want to marry women with a child during a recession because they cannot afford child-care costs, the marriage rate of women with children will fall when the unemployment rate is high. In this case, the female unemployment rate will not have a negative effect for older women without children. Thus, column (4) of Table 10 reports the same model estimated with women without children

[22] The year of completion is defined as the last year of enrollment taken from the Education and Training History Topical Module attached to the Wave 1 or 2. I limited the sample to women who have at least some college education because there is not much variation in age at graduation for high school graduates.

[23] Since the SIPP does not provide full employment history, I have to limit the sample to women who were unmarried and 17-35 years old at the beginning of each panel and observe their marital status only during the period of core panel survey.

only; the presence of children is not likely to be the reason for the opposite effects for the older women.

Third, older women are more likely to have left the community where they grew up and thus may have difficulty increasing search intensity effectively, especially when other women want to marry. Column (5) shows estimates from the subsample of the SIPP who have never moved from the state of birth.[24] Although the negative effect of the female unemployment rate on the marriage hazard for older women becomes weaker, it is still significantly negative. Since the positive effect for younger women is also smaller, it is hard to conclude from this result whether the negative effect comes from losing contact with the community where the woman grew up.

Overall, the differential effects across age groups are robust across different sample restrictions. It is not attributable to correlation of the age of marriage with educational background, employment status, the existence of children, or whether the woman has moved from the community where she grew up.[25] Although showing the absence of effects for these factors is not direct evidence, it suggests that the effects of gender specific unemployment rates are varying with the age of the woman herself, rather than with another factor correlated with age.

6 Interpretation in a Search Theoretic Framework

The probability of marriage is the product of the probability that a woman meets a man and the probability that they agree to marry conditional on meeting. In standard search models, the probability that a woman meets a man is called the arrival rate and often depends on search effort, and the reservation quality of an acceptable match determines whether they agree to marry conditional on meeting. Therefore, the increase in marriage incidence for young women due to worse female labor market conditions must come either through an increase in search intensity for mates or a decrease in the reservation match quality that is compensated for by a temporary rise in the relative earnings of men. While lowering the reservation match quality is expected to

[24] Although women with college education are more likely to have left their state of birth, Table 3 shows that the female unemployment rate has a negative effect on marriage incidence for older women regardless of their educational backgrounds.

[25] In addition, Appendix section B.3 and Table B.4 show that even if the benefit from legal marriage could vary with age, the estimated effect of the female unemployment rate does not significantly change by splitting the sample by the existence of legal institutions that increases the benefit from being legally married.

increase the probability of divorce after the recovery of female labor market conditions, increased search effort accelerates the timing of marriage without affecting the likelihood of divorce in future. Hence, the empirical results presented in section 4.1 can be interpreted as evidence for a change in search intensity rather than a change in reservation match quality.

In a standard dynamic search model of marriage formation such as Mortensen (1988), the gain from marriage for a woman is typically defined as follows: $gain = \theta^w(V_M - V_S^m - V_S^w)$, where the woman's share of the total gain from marriage is given by a constant $\theta^w \in (0, 1)$, and V_M, V_S^m and V_S^w are the value of marriage and the values of the single state for a man and for a woman. Assuming that an increase in the gain from marriage causes an increase in the actual marriage incidence, the positive effect of the female unemployment rate on marriage incidence implies that *gain* is increasing in the female unemployment rate. Likewise, the negative effect of the male-female gap implies that *gain* is decreasing in the male-female gap.

In order to build a full model which incorporates the effects of labor market conditions, however, I would have to specify the value functions which depend not only on male and female unemployment rates and the match quality but also on the stock of men and women who are looking for matches and on a matching function which generates the arrival rate from both endogenous search intensity and the supply of men and women in each period. It is however possible to see some key insights from a reduced form model in which gender-specific labor market conditions shift the demand for marriage with a given level of match quality.

For simplification, the net effects of female and male unemployment rates are summarized as "labor market conditions" U_t, which increases *gain*. The gain from a particular marriage also depends on the match quality, x, and the woman's preference for marriage, π.[26] Thus, under the assumption of transferable utility between spouses,[27] a women and a man agree to marry if and only if $gain(x, \pi, U_t) > 0$. The "demand" for marriage, $D_t(x, U_t)$, is thus defined as the number of women whose π satisfies $gain(x, \pi, U_t) > 0$ with the given (x, U_t). In other words, $D_t(x, U_t)$ is

[26] Since the model is described from the viewpoint of women, men's preference for marriage, which also affects *gain*, is included in the match quality x.

[27] Utility is assumed to be transferable between spouses. This assumption implies that a single woman is willing to transfer utility to the man whom she meets up to the point where the man is also better off from the marriage or she is no longer better off from the marriage. Hence, the benefit from marriage for a woman and that for a man can reduced into one parameter which represents the benefit from marriage for the woman net of the transfer to the man.

the number of women who is willing to marry with match quality x under labor market condition U_t.

For further simplification, assume $\partial gain/\partial x \geq 0$, $\partial gain/\partial \pi \geq 0$, and $\partial gain/\partial U_t \geq 0$ for any (x, π, U_t). Then, the demand curve of marriage increases in x for any given U_t. On the other hand, taking the distribution of x and the number of men and women who are single as given, the supply of matches whose qualities are greater than x decreases in x. Figure 6 shows the demand curve $D_t(x, U_t)$ and the supply curve $S_t(x)$, taking the number of women willing to marry and the number of available matches on the horizontal axis and the match quality on the vertical axis. The intersection of $D_t(x, U_t)$ and $S_t(x)$ determines the equilibrium number of marriage formed in this period and the reservation match quality, i.e. the lower bound of accepted matches.

Figure 6 also illustrates how U_t affects the number of marriage. An increase in U from U_0 to U_1 shifts the demand curve to the right. In Case 1, the supply curve does not shift because the number of men and women and the probability that they meet are exogenously given. Thus, the intersection moves so that the number of marriages increases from M_0 to M_1 and the reservation match quality falls from x_0^* to x_1^*. In this case, the extra marriages induced by gender specific labor market shocks are on average poorer matches and more likely to end in divorce than marriages formed without such shocks.

Yet, the number of men whom a woman meets can be increased by putting in more search effort, as long as there are some frictions in the marriage market.[28] Search effort reduces these frictions and increases the probability that a woman meets a man. From each woman's point of view, this increase in the arrival rate can be interpreted as an increase in the number of available matches or the supply of matches. Presumably, when marriage is more attractive for women, they will put in more search effort. Therefore, the probability of meeting, or the arrival rate, is likely to increase in U_t. Case 2 of Figure 1 illustrates this situation: the change from U_0 to U_1 shifts the demand curve to the right, but the change of the arrival rate from λ_0 to λ_1 shifts the supply

[28] The probability of meeting a new potential mate increases not only when women who are already searching for mates intensify the search, but also when women who had not begun searching yet choose to start searching. In addition, a typical couple spends a few years seeing each other before getting married, which Becker et al (1977) call "intensive" search to figure out whether the current partner is good enough to marry. I include effort to shorten this intensive search process in what I call "search effort" here, as well as "extensive" search effort to meet a new potential mate.

curve to the right, too. As a consequence, despite the increased number of marriage from M_0 to M_1, the reservation quality does not necessarily fall.

Although I do not specify how λ is determined, it is worth emphasizing that search effort and the reservation match quality are simultaneously determined. Therefore, the extent to which women are willing to lower their reservation match quality depends on the relative cost of increasing the arrival rate by putting in more search effort. If lowering the match quality is much more costly than increasing search efforts, it possible for women to decide not to lower the reservation match quality. Case 2 of Figure 6 describes this case; search intensity increases up to the point where the marginal cost of additional effort equals to the marginal benefit from it, holding the reservation match quality unchanged.

The empirical results presented in the previous section imply that the match quality of newly formed marriage does not change in response to labor market conditions. Hence, Case 2 is more likely to be the case. Also, if the reservation quality does not change, an increase in marriage incidence due to a temporary increase in search effort may not necessarily raise the fraction who eventually marry in the cohort much, as long as the total number of matches with $x \geq x^*$ is limited.

Furthermore, the limited supply of men implies that each woman's search effort creates a negative externality on other women's marriage incidence.[29] This negative externality may be able to explain the negative effect of the female unemployment rate on marriage incidence for older women. Older women may be crowded out by the increased search effort by younger women for several reasons. First, men may prefer to marry with younger women because they simply appreciate youth or they believe women who have not married until the late twenties are negatively selected. Second, women who hope to marry by a certain age have already put in much search effort, and they may not have room to adjust the timing of marriage according to labor market conditions.

If the crowding-out by young women's search effort is the main reason for the opposite effects on marriage incidence for older women, the increase in the number of young women getting

[29] This negative externality can be created not only by direct competitions in the "extensive" search process but also by better outside options for men which prolong the "intensive" search process necessary to reach an agreement to marry.

married must exceed the decrease in the number older women getting married. The overall effects shown in Table 3A are consistent with this prediction. Yet, if young women's search efforts crowd out older women, the reservation quality of husband for older women should fall when marriage incidence for younger women increases. The results presented in Appendix Tables B2 are not quite consistent with this prediction although it may be because older women are not willing to lower their reservation match quality in order to exploit the temporary increase in the gain from marriage due to labor market shocks.

An alternative explanation is that, if women are inherently different in terms of taste for marriage, and if women who choose not to marry by the mid-twenties prefer to marry when the female unemployment rate is low, the estimated coefficients are picking up heterogeneous inherent taste for marriage. It could also be the case that men become more selective when the female unemployment rate is high for some reason and choose not to marry women who remain unmarried because men believe these women are negatively selected. I cannot rule out such possibilities, however, the crowding out effect is at least one of the possible explanations for the negative effect of the female unemployment rate on marriage incidence for older women.

7 Concluding Remarks

The increase in marriage incidence due to relatively worse female labor market conditions is driven by the acceleration of marriage timing by women in their teens and early twenties. Marrying younger in response to labor market conditions does not increase the probability of future divorce. Moreover, this increased marriage incidence at the teens and early twenties does not increase the fraction of women who will ever marry in the long run. That is, the response of the marriage rate of young women to gender-specific unemployment rates is a timing effect for those who would marry anyway, rather than a permanent increase in the number of women who ever marry. Furthermore, despite the significant effects on marriage incidence, gender-specific labor market conditions in youth do not significantly affect the fertility and labor market outcomes observed in the mid-thirties, except for the subtle change in weeks worked.

These results cast doubt on the view that a further improvement of women's status in the labor market will lead to a further decline in the marriage rate. It is true that women delay marriage to exploit better labor market opportunities and it lowers the marriage rate in the population. However, women who would marry eventually do marry anyway, and there seems to be no effects on marriage stability. Moreover, a decrease in marriage incidence due to labor market fluctuations does not lead to an increase in single parenthood, in contrast to the prevailed view that changes in marriage incidence imply changes in the number of women who ever enter stable marriage.

Yet, it must be kept in mind that this paper has estimated the effects of temporary fluctuations in gender specific labor market conditions, which, by definition, do not affect the permanent income or long-run labor market prospects. Permanent shifts in labor market prospects of women may change the perceived value of marriage in the long run, although it is difficult to distinguish such permanent shifts from other changes in trends. Nonetheless, this study provides at least some evidence that the response of contemporaneous marriage incidence is basically an inter-temporal substitution behavior by showing the absence of the long-run effects.

A Data Appendix

A.1 SIPP

Wave 2 of the SIPP contains retrospective information of marriage history of up to three marriages, the dates of birth of the first and last children and limited migration history, as well as basic demographic information as of the date of survey. I format the dataset as if it were a set of panel surveys interviewed on January 1st every year since 1978. When I estimate the effects of contemporaneous unemployment rates (e.g. Table 3), the dependent variable is defined as the probability of getting married during the year when the woman becomes age x.

Variables in the SIPP Core Panels are collected either on a monthly basis or once in each wave. The dataset is collapsed to an annual basis by taking either the value at January or the average over the year of each variable and merged with the variables from the Wave 2 Topical Modules. Consequently, the sample is restricted to those who were present in the household at both the Wave 2 interview and the first interview in the corresponding calendar year. Since some variables (e.g. spouse's age, the number of children, employment status) are available only from the Core Panel data, I have to restrict the sample for analysis using such variables accordingly. Table A1 shows summary statistics for each subsample.

The Migration History Topical Module includes information on: the state of residence on the date of survey; the year and month when the respondents moved to this state; the state of previous residence (if there is any), which can be the same state as the current residence; the year and month when the respondents moved into current and previous residences; and the state or country of birth. Thus, the state of residence can be retrieved back to the earlier of the dates moving to the current state or moving to the previous residence. I also assume that those whose state of the previous residence is the same as the state of birth had lived in that state since their birth until they moved to the state of current residence. Table A2 shows the fraction of the observations in the base sample whose state of residence in the year is identified. The state of residence at marriage is determined for 75.7 percent (77.9 percent for still in the first marriage, 69.2 percent for divorced) of all first marriages in the sample.

A.2 PSID

Extracting marriage history from the PSID is tricky because for most couples, one of the spouses had not been included in the survey until they married. Also, although the PSID tries to follow every member of the households interviewed in the previous year, many people (especially those who joined in the sample at their marriage) disappear when they divorce. Therefore, for the marriage hazard regression (Table 3), I restricted the sample to women who were in the original sample in 1968 and their daughters so that newly joined cohabiting partners are not included in the sample.

For the analysis of the effects of gender-specific unemployment rates at marriage on spouse's characteristics (Table 5) and the number of children (Table7), the dataset is constructed in the following way:

1. Extract all first marriages of women in Marriage History File 1985-2005.
2. Merge it with men in the same file, and fill the dates of marriage and divorce missing from wives' records with their husbands'. This mitigates the attrition due to divorce to some

extent. Now the sample size is equal to the number of marriages.

3. Merge the extracted marriage history data with a longitudinal file of all women from the main PSID, which includes all non interviews, and then merge it with a longitudinal file of all men, which also includes all non interviews.

This yields a panel dataset of married couples containing all available information of both wives and husbands.

The year of birth, the years of schooling and the number of children are taken from Individual Data File 1968-2005. Religion is taken from Main Family Data files 1985-2005 and available only for household heads and their wives. Since both no religion and no answer seemed to be coded as "0", I assume the respondent has no religion if he/she has answered 0 for all interviews, and replace the others with the non-zero answers by the same person in other years. To make the coding consistent over time, religions are recoded for all years as follows: no religion, Roman Catholic, Jewish, Baptist, Lutheran, Methodist/African Methodist, Presbyterian, Episcopalian, Protestant unspecified, Greek/Russian/Eastern orthodox, and other religions. While most people do not change their religion, for those who have changed religion since marriage, the religion in the next year of marriage is used to define the variable "same religion".

Race is available only for respondents who have been a head or a wife of an interviewed household at some point in 1968-2005 or children of mothers included in Fertility History File 1985-2005. For those whose race is not determined directly, if they are born to a family whose head and wife are of the same race, I assume they are also of the same race.

B Robustness Checks

B.1 Instrumenting for Unemployment Rates

Following Bartik (1991) and Blanchard and Katz (1992), I constructed instruments for state unemployment rates by gender by taking weighted average of nation-wide unemployment rates by industry-occupation cells using the industry-occupation composition of each state and gender as the weights. I use industry-occupation specific unemployment rates instead of log employment changes used by Bartik (1991) and Blanchard and Katz (1992), because log employment changes might pick up differences in changes in marriage and labor supply decisions of women with different skill levels, which might produce a persistent but not permanent effect on the marriage rate in each state. Another important difference from Bartik (1991) and Blanchard and Katz (1992) is that weights are fixed over the entire period, rather than based on the industry-occupation composition in the previous year.

28 industry-occupation cells are defined as follows. First, based on the one-digit occupation code in the CPS, I split all jobs into two groups: white collar (WC), and blue collar, service and farmers (BC). Next, I recode the major industry code so that the coding is consistent over time and each industry has a large enough sample size: agriculture/forestry/fishery, mining, construction, manufacturing-durable, manufacturing-nondurable, transportation/communication/utility, wholesale trade, retail trade, finance, personal service (including private household), business and repair service, entertainment service, hospitals and other medical service, educational service, other service, and public sector. Then, I split each industry into BC and WC, except for the following four industries whose sample size of BC or WC is too small: agriculture/forestry/fishery, mining, finance, and entertainment service.

The ideal instruments for gender-specific state unemployment rates should affect labor demand for men and for women differently and be independent from marriage behaviors of the population. Also, they must vary across states over years so that state- and year- fixed effects can be controlled for. As pointed by Blanchard and Katz (1992), the weighted average of industry-occupation specific unemployment rates is a valid instrument for the state unemployment rate if national-level changes in unemployment rates by industry-occupation are not correlated with labor supply shocks in the state. Panel A of Table B1 shows that the first stage correlation is strong enough even after controlling for state- and year- fixed effects. However, Panel B shows that much of the variation in the instruments is absorbed by these fixed effects. Therefore, we should keep in mind that small noises might affect the IV estimates substantially.

To make the estimation feasible, the Cox's proportional hazard model in equation (1) is modified into the following linear probability model:

$$M_{its} = \alpha_{age} u_{ts}^{w} + \beta_{age}(u_{ts}^{h} - u_{ts}^{w}) + age_i + \eta_s + \xi_t + \varepsilon_{its} \tag{7}$$

where age_i and ξ_t are dummy variables for each single year age and calendar year. Appendix Table B1 presents the estimated α and β in (7). The first column shows the OLS estimates. The point estimates imply that 1 percentage-point rise in the female unemployment rates increases the marriage hazard for 17-20 years old women by 0.2 percentage point, which is much smaller than the effect estimated by the Cox's proportional hazard model shown in Figure 3. Nevertheless, the signs and the relative size of coefficients of different age categories remain the same. The second column reports the 2SLS estimators. The point estimates are in the range between the effects estimated by the OLS and by the Cox hazard model.

B.2 The Effects by Age at Marriage

Since the effects of contemporaneous unemployment rates by gender on the marriage hazard vary with women's age, the effects of unemployment rates at marriage could also vary with women's age at marriage. Thus, I allow the coefficients of unemployment rates at marriage to vary with the age at marriage in the same way as in Table 4B. Appendix Table B2 presents the effects on spouses' characteristics. The results are noisy but similar to Table 5. Appendix Table B3 shows the effects on the number of children. It confirms that the effects are insignificant, as shown in Table 7.

B.3 Legal Institutions

Benefits from being legally married such as spousal benefits from employers or health insurance coverage may have different importance for women of different ages. If so, the availability of such benefits could change the effects of gender-specific unemployment rates on marriage incidence. Since it is hard to obtain individual-level information on this, I use indicators of state laws that give benefits to legally married wives as proxies for differences between legal marriage and cohabitation. Specifically, I use the following two indicators. One is whether the state's marital-property law follows the common law rule, under which the distribution of property upon divorce is directed toward husbands.[30] Since the financial gain from having a high earning spouse is lower in such states, worse female labor market conditions should have less positive effects on marriage incidence in states with the common law property right rule. The other indicator is whether the state of residence recognizes common law marriages; in the states where the common law marriages are recognized, financial benefits from legal marriage are supposed to be smaller than in other states, therefore the coefficients of gender-specific unemployment rates will be closer to zero.[31]

Appendix Table B4 presents the estimated coefficients of gender-specific unemployment rates in equation (1). The difference in the rule on the property division at divorce makes no difference in the effects of labor market conditions on marriage incidence. Moreover, contrary to expectations, the coefficient of the female unemployment rate is more negative for older women in the states where common law marriages are recognized. Therefore, the differential benefits from being legally married do not seem to be able to explain why the effects of unemployment rates on marriage incidence vary with women's age.

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[30] The classification is based on Gray (1998). States with the common law property rule are Alabama, Florida, Georgia, Maryland, Massachusetts, Mississippi, Missouri, Montana, New York, North Carolina, Ohio, Pennsylvania, South Carolina, Tennessee, Virginia, and West Virginia.

[31] States that recognize the common law marriages are the following: Alabama, Colorado, District of Columbia, Georgia (until 1/1/97), Idaho (until 1/1/96), Iowa, Kansas, Montana, New Hampshire (for inheritance purposes only), Ohio (until 10/10/91), Oklahoma, Pennsylvania (until 1/1/05), Rhode Island, South Carolina, Texas, Utah.

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Table 1 Summary Statistics

Non-Hispanic white women who had not married until age 16 or 1978

		(A) SIPP	(B) PSID
Year of birth:	mean	1965.4	1967.3
	min	1956	1956
	max	1980	1980
Year of 1st marriage:	mean	1986.0	1989.1
	min	1978	1978
	max	2001	2005
Average age at first marriage		22.4	23.0
Average age at birth of 1st child		23.4	23.5
Sample size (persons)		52,217	3,698
% by schooling	Dropouts	9.2%	9.2%
	High school	30.5%	34.6%
	Some college	35.1%	28.0%
	College	25.3%	28.2%

Note: The statistics of the first marriage are based on those who married by the age of 35, and those of birth of the first child are based on those who had a child by their last interview. The SIPP sample is weighted by the sample weights.

Table 2: Summary statistics of gender-specific unemployment rates, non-Hispanic white 16-40 years old, 45 states, 1978-2003.

	Mean	Standard Deviation	p75-p25
Female unemployment rate	6.61	2.12	2.75
residuals	-	1.04	1.27
Male unemployment rate	6.64	2.51	2.93
Male- Female gap	0.03	1.44	1.72
residuals	-	0.97	1.24
Number of observations	1,170		

Note: The residuals are net of the other gender's unemployment rate and year- and state- fixed effects.

Table 3: Effects of the female unemployment rate and the male-female gap in the unemployment rate on the marriage hazard
(Cox's proportional hazard model)

A. Pooled

	SIPP				
	All women	High school Or less	Some college	BA or more (>=21 yrs old)	PSID
Female unemployment rate	0.048***	0.022	0.029**	0.069***	0.029
	[0.012]	[0.014]	[0.014]	[0.014]	[0.027]
Male-female gap in u rates	-0.036***	-0.033**	-0.028**	-0.014	-0.009
	[0.011]	[0.014]	[0.014]	[0.015]	[0.028]
Observations	269,621	89,812	94,513	58,249	11,109
Persons	39,949	17,240	14,333	10,135	1,832

B. By age

	SIPP				
	All women	High school or less	Some college	BA or more (>=21 yrs old)	PSID
Female unemployment rate* woman's age					
17-20 years old	0.129***	0.064***	0.115***	--	0.117***
	[0.014]	[0.015]	[0.016]		[0.034]
21-23 years old	0.046***	-0.004	0.020	0.170***	-0.006
	[0.011]	[0.014]	[0.014]	[0.014]	[0.028]
24-27 years old	-0.070***	-0.088***	-0.080***	0.017	-0.106***
	[0.014]	[0.020]	[0.016]	[0.012]	[0.031]
28 or older	-0.217***	-0.228***	-0.218***	-0.147***	-0.271***
	[0.030]	[0.037]	[0.029]	[0.026]	[0.044]
Male-female gap* woman's age					
17-20 years old	-0.097***	-0.070***	-0.103***	--	-0.170***
	[0.024]	[0.021]	[0.027]		[0.044]
21-23 years old	-0.011	-0.015	0.029	-0.065***	0.047
	[0.012]	[0.015]	[0.019]	[0.025]	[0.038]
24-27 years old	0.038**	0.079***	0.006	0.020	0.127**
	[0.016]	[0.021]	[0.025]	[0.019]	[0.057]
28 or older	-0.007	-0.021	-0.033	0.006	0.198**
	[0.038]	[0.053]	[0.054]	[0.026]	[0.089]
Observations	269,621	89,812	94,513	58,249	11,109
Persons	39,949	17,240	14,333	10,135	1,832

Note: Standard errors in brackets are clustered by the state of residence. The baseline hazard depends on age and is stratified by the year of birth. All columns include controls for state fixed effects.

Table 4: Effects of the past gender-specific unemployment rates on divorce (SIPP)

A. Pooled effects on the probability of divorce within 5 or 10 years (Probit, marginal effects)

Dependent variable:	5-year divorce rate		10-year divorce rate	
Unemployment rates at:	Marriage	Age 18-20	Marriage	Age 18-20
Female unemployment rate	-0.001	0.001	0.005	0.004
	[0.003]	[0.003]	[0.005]	[0.006]
Male-female gap in unemp. rates	-0.001	0.002	0.001	0.023***
	[0.003]	[0.003]	[0.005]	[0.005]
Observations	19,997	19,544	13,184	12,155

B. Effects by wife's age at marriage on the probability of divorce within 5 years
(Probit, marginal effects)

Dependent variable:	5-year divorce rate		10-year divorce rate	
Unemployment rates at:	Marriage	Age 18-20	Marriage	Age 18-20
Female unemployment rate * wife's age at marriage				
20 years old or younger (α_1)	-0.003	-0.004	0.001	-0.001
	[0.003]	[0.004]	[0.006]	[0.007]
21-23 years old (α_2)	0.003	0.003	0.009*	0.01
	[0.003]	[0.004]	[0.005]	[0.007]
24-27 years old (α_3)	-0.002	0.001	0.007	-0.001
	[0.004]	[0.005]	[0.006]	[0.010]
28 years old or older (α_4)	-0.002	0.009*	0.007	0.018
	[0.004]	[0.005]	[0.009]	[0.012]
Male-female gap in unemployment rates * wife's age at marriage				
20 years old or younger (β_1)	0.002	0.004	0.002	0.028***
	[0.003]	[0.004]	[0.005]	[0.006]
21-23 years old (β_2)	-0.004	0.001	0.000	0.019***
	[0.004]	[0.005]	[0.006]	[0.006]
24-27 years old (β_3)	-0.001	0.007	0.006	0.031***
	[0.004]	[0.004]	[0.007]	[0.009]
28 years old or older (β_4)	0.008	-0.001	0.016	0.002
	[0.008]	[0.007]	[0.015]	[0.016]
Observations	19,997	19,544	13,184	12,155
Test stats for α_1=α_2=α_3=α_4=0	5.98	10.79	5.64	8.26
(P-value)	(0.201)	(0.029)	(0.228)	(0.082)
Test stats for β_1=β_2=β_3=β_4=0	4.80	3.69	1.73	24.43
(P-value)	(0.309)	(0.449)	(0.785)	(0.000)

Note: Standard errors in brackets are clustered by the state of residence at marriage/birth. Marginal effects are evaluated at the mean of X. Controls included in the regressions but omitted from the table are dummy variables for the year of marriage/birth and for the state of marriage/birth. Panel B also includes a dummy variable for the age at marriage (21-23, 24-27, 28-35).

Table 4: Effects of the past gender-specific unemployment rates on divorce (SIPP)

C. Cox's proportional hazard model with contemporaneous unemployment rates

	pooled	by age
Contemporaneous female unemployment rate	-0.013	-0.009
	[0.012]	[0.013]
Contemporaneous male-female gap in unemp. rates	0.004	0.006
	[0.019]	[0.019]
Female unemployment rate at marriage (α_1)	0.019	0.011
	[0.016]	[0.018]
*21-23 years old (α_2)	--	0.033**
		[0.016]
*24-27 years old (α_3)	--	0.012
		[0.023]
*28 years old or older (α_4)	--	-0.003
		[0.034]
Male-female gap in unemp. rates at marriage (β_1)	0.004	0.019
	[0.021]	[0.019]
*21-23 years old (β_2)	--	-0.026
		[0.018]
*24-27 years old (β_3)	--	-0.021
		[0.026]
*28 years old or older (β_4)	--	0.040
		[0.056]
Dummy variable for wife's age at marriage		
21-23 years old	--	-0.835***
		[0.128]
24-27 years old	--	-1.024***
		[0.188]
28 or older	--	-1.094***
		[0.236]
Observations	219,075	219,075
Persons	24,117	24,117
Test stats for α_1=α_2=α_3=α_4=0	--	6.65
(P-value)		(0.156)
Test stats for β_1=β_2=β_3=β_4=0	--	5.04
(P-value)		(0.283)

Note: Standard errors in brackets are clustered by state of residence at marriage. Baseline hazard is depends on years since marriage and is stratified by the year of marriage. Both columns include controls for state-of-marriage fixed effects.

Table 5: Gender specific unemployment rates at marriage and spouses' characteristics
Liner OLS for (1) (2) / Probit (marginal effects) for (3) (4)

A. unemployment rates at marriage

Variables	(1) Husband's age-wife's age		(2) Husband's years of schooling		(3) Same religion	(4) Same race
Dataset	SIPP	PSID	SIPP	PSID	PSID	PSID
Female unemployment rate	-0.107	0.199**	-0.061*	0.049	-0.025*	-0.010*
at marriage	[0.097]	[0.095]	[0.034]	[0.038]	[0.014]	[0.005]
Male-female gap in unemp. rates	-0.071	-0.222***	-0.021	0.008	-0.025	0.007
at marriage	[0.093]	[0.082]	[0.048]	[0.050]	[0.016]	[0.005]
Wife's years of schooling			0.612***	0.562***		
			[0.018]	[0.038]		
Observations	4,192	1,622	4,112	1,498	1,430	1,081

B. Unemployment rates at age 18-20 (SIPP only)

Variables	(1) Husband's age-wife's age	(2) Husband's years of schooling
Female unemployment rate	-0.020	0.010
at marriage	[0.087]	[0.035]
Male-female gap in unemp. rates	-0.075	-0.010
at marriage	[0.124]	[0.059]
Wife's years of schooling		0.600***
		[0.014]
Observations	4,488	4,408

Note: Standard errors in brackets are clustered by the state of residence at marriage for Panel A and the state of birth for Panel B. Controls included in the regressions but omitted from the table are dummy variables for the year of marriage/birth and for the state of marriage/birth.

Table 6: Effects of gender-specific unemployment rates that the cohort faced at fixed ages on the fraction of women who have ever married in the cohort at different ages.
(Birth cohorts 1960-70 in SIPP 2001 and SIPP 2004)

Age	Median age of marriage	22	24	26	28	30	35
Female unemp. rate	-0.263***	0.019***	0.019***	0.015**	0.007	0.007	0.000
at age 18-20	[0.027]	[0.007]	[0.006]	[0.006]	[0.006]	[0.006]	[0.005]
Male – female u. rate	0.131***	-0.008	-0.012	-0.009	-0.008	0.001	0.000
At age 18-20	[0.035]	[0.008]	[0.008]	[0.009]	[0.007]	[0.007]	[0.007]
Observations	9,536	486	486	486	486	486	397

Note: The first column is the median regression at the individual level, and the other columns are linear regressions at the cohort level (separate regressions by age) weighted by the sum of SIPP sample weights of all observations in each cell. Standard errors in brackets are clustered by the state of birth. Controls included in the regressions but omitted from the table are dummy variables for the year of birth and for the state of birth.

Table 7: Effects of gender-specific unemployment rates at the time of marriage on the number of children

Sample and dependent variable:	SIPP: the number of children at survey for women who were 35-37 years old at survey		PSID: the number of children at age 35
Unemployment rates at:	Marriage	Age 18-20	Marriage
Female unemployment rate	-0.015	-0.025	0.007
	[0.020]	[0.027]	[0.031]
Male-female gap in unemp. rates	-0.036	0.002	0.023
	[0.024]	[0.039]	[0.039]
Observations	2,806	2621	946

Note: Standard errors in brackets are clustered by the state of residence at marriage/birth. Controls included in the regressions but omitted from the table are dummy variables for the year of marriage/birth and for the state of marriage/birth.

Table 8: Effects of gender-specific unemployment rates that the cohort faced at fixed ages on the fraction of women who have ever had a child in the cohort at different ages.
(Birth cohorts 1960-70 in SIPP 2001 and SIPP 2004)

	Median age at first births	Have a child by age 25	Have a child by age 30	Have a child by age 35	Have a child before marriage	Ever been a single mother
Female unemp. rate	-0.027	0.002	0.002	0.014	-0.005	0.003
at age 18-20	[0.061]	[0.008]	[0.007]	[0.011]	[0.005]	[0.007]
Male – female u. rates	0.489***	-0.014	-0.020**	-0.005	0.000	0.001
At age 18-20	[0.080]	[0.010]	[0.008]	[0.011]	[0.006]	[0.007]
Observations	9536	486	486	486	486	486

Note: The first column is the median regression at the individual level, and the other columns are linear regressions at the cohort level (separate regressions by age) weighted by the sum of SIPP sample weights of all observations in each cell. Standard errors in brackets are clustered by the state of birth. Controls included in the regressions but omitted from the table are dummy variables for the year of birth and for the state of birth.

Table 9: Effects of gender-specific unemployment rates that a woman faced at age 18-20 on various outcomes observed at age 33-37

A. All women

	Log real personal earnings	Log real household income	Weeks worked last year (% in all weeks)	Full time employment	Pr (ever married)
Female unemp. rate	0.011	-0.005	0.012**	0.008	-0.001
At age 18-20	[0.019]	[0.008]	[0.006]	[0.007]	[0.006]
Male – female u. rates	0.021	-0.003	0.002	0.006	0.006
At age 18-20	[0.020]	[0.011]	[0.008]	[0.008]	[0.007]
Observations	10,160	12,726	12,746	12,746	15,098

B. Women who have married

	Log real personal earnings of the woman	Log real household income	Weeks worked last year (% in all weeks)	Full time employment	Log real personal earnings of the husband
Female unemp. rate	0.012	0.000	0.017**	0.01	-0.012
at age 18-20	[0.022]	[0.008]	[0.007]	[0.007]	[0.014]
Male – female u. rates	0.025	-0.012	0.003	0.006	-0.018
At age 18-20	[0.020]	[0.013]	[0.009]	[0.008]	[0.016]
Observations	8,877	11,263	11,276	11,276	9,077

Note: Standard errors in brackets are clustered by the state of birth. Controls included in the regressions but omitted from the table are dummy variables for educational backgrounds (high school graduate, some college, BA or more), for current age, for the year of birth and for the state of birth.

Table 10: marriage hazard regressions splitting sample by potentially related factors

Cox's proportional hazard model
(specification identical to Table 3b estimated with different subsamples)

	(1) Baseline (same as Column 1 of Table 3)	(2) Schl>12, age >=21, within 3 yrs since graduation	(3) Not employed fulltime and not enrolled fulltime in the previous year (core panel)	(4) Not having a child	(5) Living in the state of birth
Female unemployment rate* woman's age					
17-20 years old	0.129***	0.197***	0.391***	0.125***	0.065***
	[0.014]	[0.023]	[0.098]	[0.016]	[0.014]
21-23 years old	0.046***	0.052***	0.154*	0.046***	0.022**
	[0.011]	[0.016]	[0.088]	[0.012]	[0.010]
24-27 years old	-0.070***	-0.103***	-0.056	-0.073***	-0.042***
	[0.014]	[0.021]	[0.112]	[0.014]	[0.015]
28 or older	-0.217***	-0.285***	-0.134	-0.223***	-0.105***
	[0.030]	[0.057]	[0.129]	[0.027]	[0.029]
Male –female gap in unemp. rate* woman's age					
17-20 years old	-0.097***	-0.099**	-0.139	-0.098***	-0.078***
	[0.024]	[0.046]	[0.110]	[0.028]	[0.020]
21-23 years old	-0.011	-0.008	0.105	-0.009	0.008
	[0.012]	[0.024]	[0.091]	[0.014]	[0.015]
24-27 years old	0.038**	0.023	0.101	0.043***	0.054***
	[0.016]	[0.033]	[0.138]	[0.016]	[0.018]
28 or older	-0.007	0.018	-0.015	0.016	0.022
	[0.038]	[0.056]	[0.195]	[0.039]	[0.036]
Observations	269,621	30,025	8,209	234,716	185,750
Persons	39,949	9,540	5,097	37,673	26,503

Note: Standard errors in brackets are clustered by the state of residence. The baseline hazard depends on age and is stratified by the year of birth. All columns include controls for state fixed effects.

Figure 1: % of women who have ever married, by age
non-Hispanic white women born in 1956-80; SIPP 90-04 (weighted)

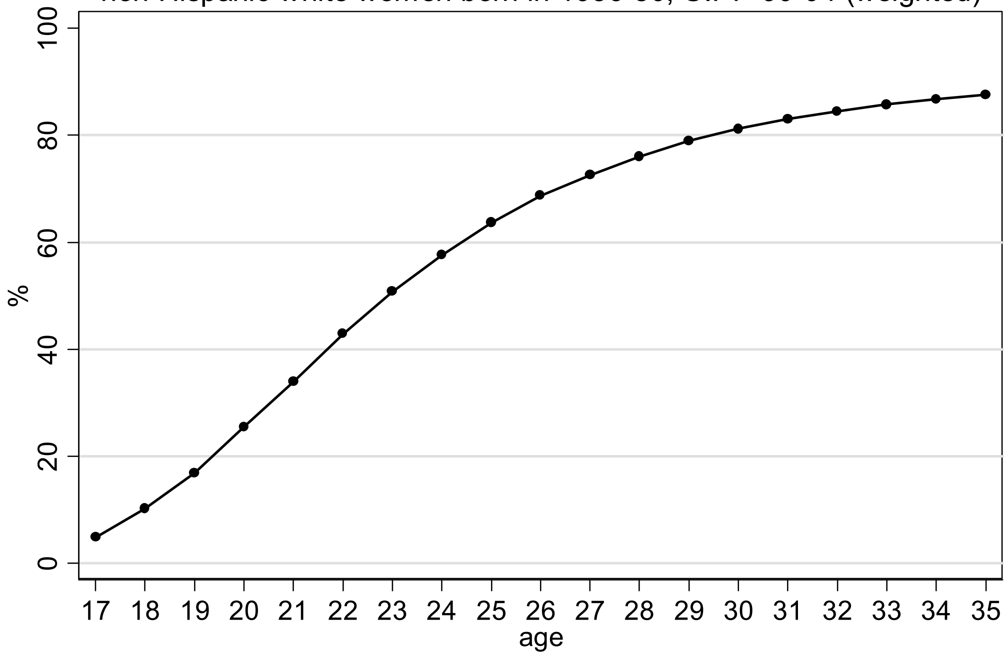


Figure 2: Female unemployment rate and male-female gap for selected states (Non-Hispanic whites, age 16-40, 1978-2003, CPS)

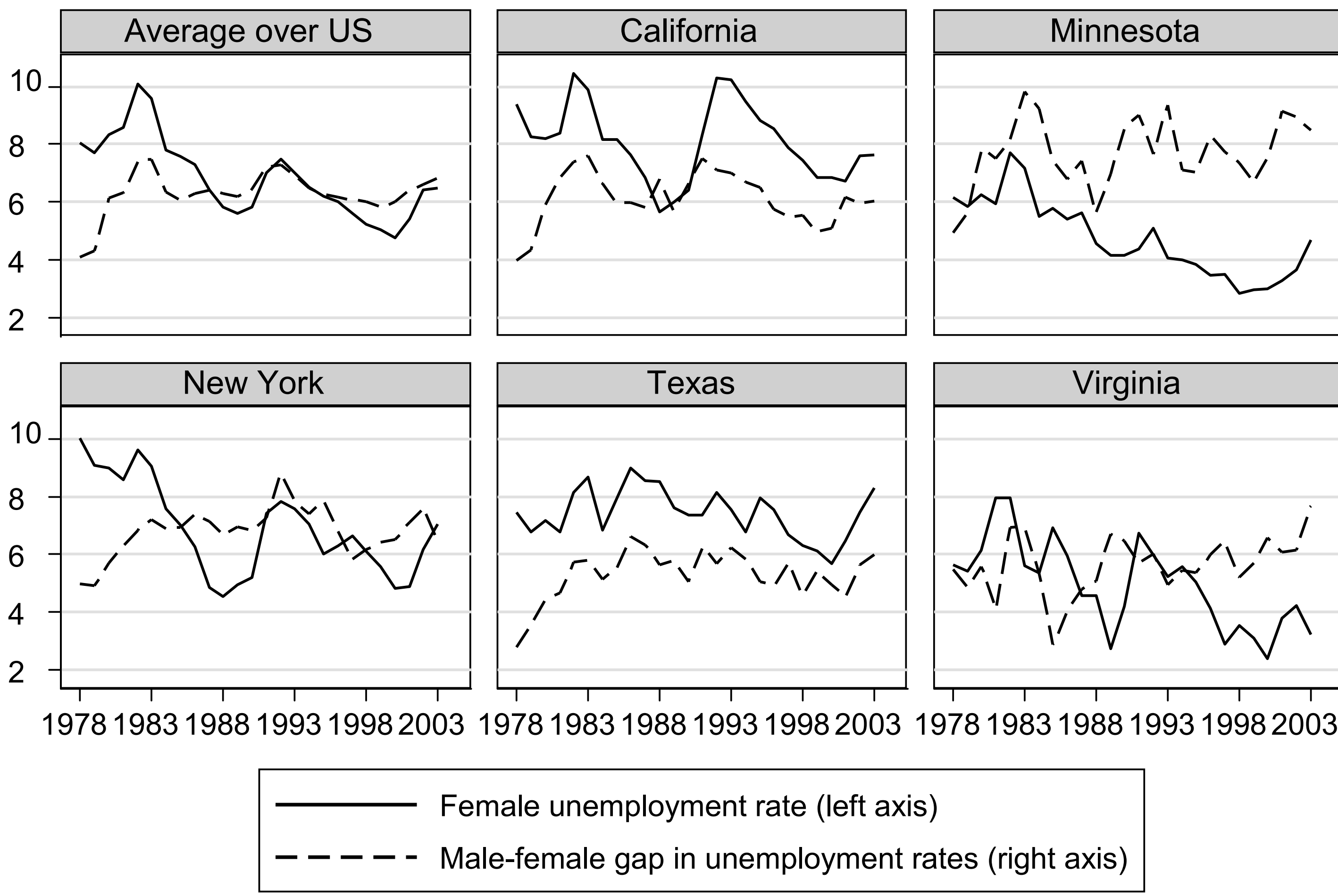


Figure 3: Female unemployment rate in youth and various outcomes at the cohort level (without controlls for year and state fixed effects)

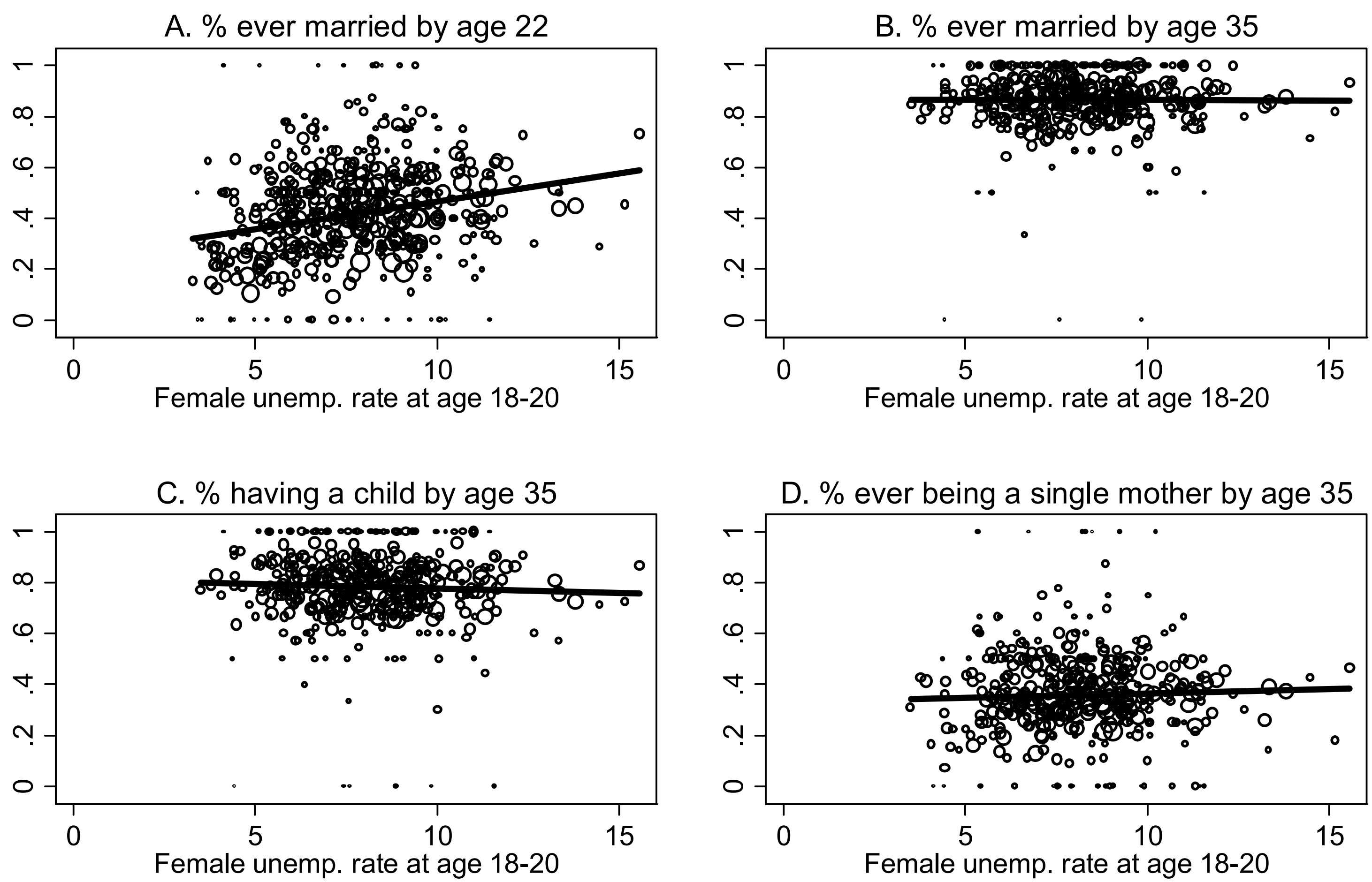


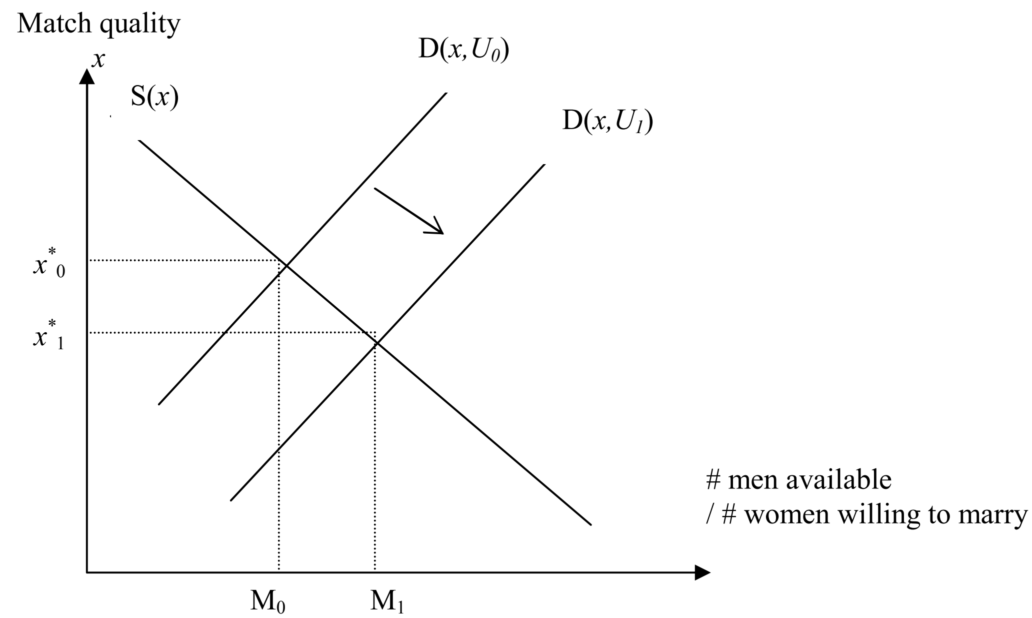
Figure 4: Effects of 1 %-point rise in gender-specific unemployment rates on age-specific marriage hazard (in %-point)

Note: (Exp(coefficient)-1)*empirical hazard by age. Coefficients in the Cox's proportional hazard model are estimated with the SIPP. Dotted lines show 95%-confidence intervals.

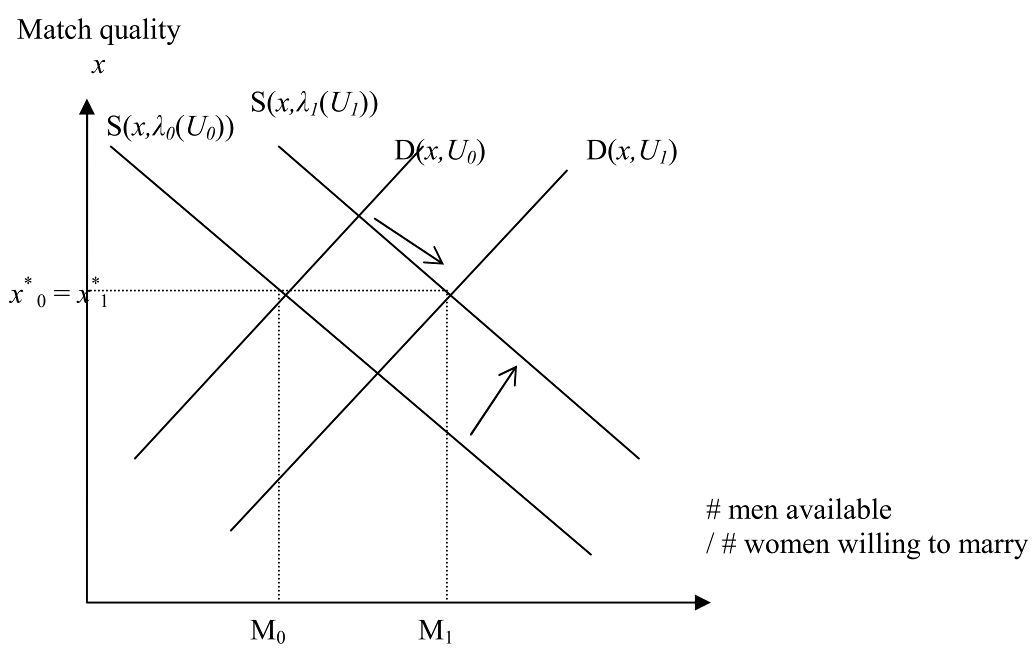
Figure 5: % ever married by each age and the female unemployment rate experienced at age 18-20

Figure 6: Effects of labor market conditions on marriage incidence and match quality

Case 1: No change in search intensity



Case 2: Increase in search intensity to keep the match quality unchanged



Appendix Table A1: Summary statistics of subsamples

SIPP Subsamples:

A. Women who married in 1978 ~ 5 years prior to the survey (Tables 4A 4B)

Sample size	22,234
Year of birth	1963.8
Age at marriage	22.4
5year divorce rate	14.0%
Timing of the birth of first child:	
Before marriage	11.4%
In 0-2 years	40.1%
In 3-5 years	23.3%
No child by the 5th year	25.2%

B. Couples married in 3 years prior to the survey or later and currently married (Table 5)

Sample size	4,192
Year of birth	1969.0
Wife’s age at marriage	24.9
Husband's age at marriage	27.4
Wife’s years of schooling	14.1
Husband's years of schooling	13.9

C. 34-36 years old at Wave 2 (Table 7)

Sample size	3,552
Year of birth	1963.0
% having a child	86.3%
Age at first birth (if have a child)	25.5
Number of children	1.96

D. 33-37 years old women in the core panels (Table 9)

Sample size	15,098
Year of birth	1963.8
% married	88.1%
Age at marriage (if married)	23.2
Log household income	8.3
Log person earnings (zero earnings excluded)	7.3
Husband's log person earnings (if married)	8.0
% of weeks worked in the last calendar year	75.1%
Employed full-time	57.2%

PSID: women who married in 1978 or later (Tables 5 and 7)

Sample size	1,622
Number of children at age 35	1.45
0 child	25.6%
1 child	24.1%
2 children	34.6%
3 or more	15.8%
Husband's age - wife's age	2.4
Wife’s years of schooling	13.2
Husband's years of schooling	13.2
Same religion	57.7%
Husband different race	4.7%

Appendix Table A2: % of the SIPP sample whose state of residence is determined, by calendar year and Panel.

	Panel						
Calendar year	2004	2001	1996	1993	1992	1991	1990
1978	63.0%	65.5%	65.2%	72.8%	73.5%	72.6%	75.1%
1979	63.8%	65.9%	66.6%	74.1%	74.3%	73.1%	76.7%
1980	63.9%	66.9%	67.7%	76.1%	75.6%	74.7%	78.5%
1981	65.0%	67.7%	68.3%	77.7%	76.6%	76.2%	80.2%
1982	65.4%	68.3%	69.6%	78.7%	78.0%	78.7%	81.6%
1983	66.0%	68.6%	70.6%	79.9%	79.2%	80.8%	83.1%
1984	67.0%	69.1%	71.7%	81.7%	80.1%	82.7%	84.3%
1985	67.8%	69.8%	72.8%	83.1%	82.5%	83.9%	85.6%
1986	68.7%	70.7%	73.8%	84.8%	84.8%	85.3%	88.1%
1987	69.2%	71.2%	75.3%	86.1%	86.8%	88.1%	89.4%
1988	70.3%	71.9%	76.7%	88.0%	88.8%	89.0%	91.3%
1989	71.3%	72.9%	78.3%	89.4%	89.5%	90.8%	91.6%
1990	72.5%	74.0%	80.2%	90.6%	91.7%	91.5%	
1991	73.8%	75.4%	82.1%	91.8%	92.2%		
1992	74.2%	76.2%	84.2%	92.9%			
1993	74.9%	77.0%	85.7%				
1994	75.4%	78.3%	87.6%				
1995	76.4%	79.4%	89.6%				
1996	77.4%	81.0%					
1997	78.5%	83.0%					
1998	80.0%	84.5%					
1999	81.7%	86.2%					
2000	83.8%	87.5%					
2001	86.1%						
2002	88.5%						
2003	89.8%						

Appendix Table B1: Linear 2SLS estimates of the contemporaneous effects on marriage incidence, instrumenting for state unemployment rates by gender with weighted average of nation-wide industry-occupation-gender specific unemployment rates

A. First stage (state-level regression with state- and year- fixed effects)

	Female u	Gap
Female u IV	3.792***	0.433
	[0.568]	[0.601]
Gap IV	2.746***	1.818***
	[0.430]	[0.511]
Observations	1,125	1,125

B: Summary statistics of the instruments

	Mean	Standard Deviation	p75-p25
Female u IV	5.71%	1.02%	1.40%
Residuals	--	0.08%	0.08%
Gap IV	0.29%	0.68%	0.61%
Residuals	--	0.11%	0.10%

Note: The residuals are net of year- and state- fixed effects.

C. Linear version of the marriage hazard regressions

	Linear OLS	2SLS
Female unemployment rate* woman's age		
17-20 years old	0.002**	0.006**
	[0.001]	[0.003]
21-23 years old	0.002	0.006
	[0.001]	[0.004]
24-27 years old	-0.005***	-0.001
	[0.001]	[0.004]
28 or older	-0.003	-0.006
	[0.003]	[0.006]
Male-female gap* woman's age		
17-20 years old	-0.002**	0.000
	[0.001]	[0.006]
21-23 years old	-0.002	-0.002
	[0.001]	[0.005]
24-27 years old	0.002	-0.009
	[0.002]	[0.010]
28 or older	0.003	0.001
	[0.002]	[0.016]
Observations	273,464	271,721
Persons	42,172	42,114

Note: Standard errors in brackets are clustered by the state of residence. Controls included in the regression but omitted from the table are dummy variables for the state of residence, for age, and for calendar year.

Appendix Table B2: Gender specific unemployment rates at marriage and spouses' characteristics, by age at marriage (Supplement to Table 5)

Liner OLS for (1) (2) / Probit (marginal effects) for (3) (4)

Variables	(1) Husband's age-wife's age		(2) Husband's years of schooling		(3) Same religion	(4) Same race
Dataset	SIPP	PSID	SIPP	PSID	PSID	PSID
Female unemployment rate * wife's age at marriage						
20 years old or younger (α_1)	-0.079	0.186	-0.067	0.082	-0.015	-0.013
	[0.112]	[0.113]	[0.051]	[0.057]	[0.020]	[0.009]
21-23 years old (α_2)	-0.112	0.190	-0.094*	0.081	-0.037**	-0.012**
	[0.156]	[0.117]	[0.048]	[0.051]	[0.017]	[0.006]
24-27 years old (α_3)	-0.124	0.256*	-0.011	0.005	-0.025	-0.001
	[0.104]	[0.142]	[0.048]	[0.049]	[0.018]	[0.008]
28 years old or older (α_4)	-0.096	-0.080	-0.080*	0.022	-0.010	-0.024**
	[0.127]	[0.232]	[0.041]	[0.068]	[0.027]	[0.010]
Male –female gap in unemp. rate* wife's age at marriage						
20 years old or younger (β_1)	-0.006	-0.191*	-0.124	0.018	-0.033*	0.013***
	[0.162]	[0.111]	[0.090]	[0.058]	[0.020]	[0.005]
21-23 years old (β_2)	-0.042	-0.069	-0.038	-0.107	-0.02	0.011
	[0.125]	[0.098]	[0.066]	[0.074]	[0.022]	[0.013]
24-27 years old (β_3)	-0.08	-0.214	0.042	0.034	-0.036	-0.001
	[0.095]	[0.142]	[0.045]	[0.067]	[0.024]	[0.010]
28 years old or older (β_4)	-0.101	-0.667*	-0.024	0.009	0.075**	-0.001
	[0.136]	[0.373]	[0.055]	[0.137]	[0.036]	[0.019]
Dummy variable for wife's age at marriage						
21-23 years old	-0.540	-0.662	0.471	0.427	0.141	-0.019
	[0.903]	[0.739]	[0.356]	[0.460]	[0.097]	[0.058]
24-27 years old	-0.805	-1.182	0.068	1.310***	0.024	-0.065
	[0.690]	[1.023]	[0.368]	[0.477]	[0.162]	[0.041]
28 years old or older	-1.659*	0.652	0.622	1.264*	-0.179	0.093
	[0.830]	[1.751]	[0.391]	[0.644]	[0.204]	[0.182]
Wife's years of schooling			0.586***	0.490***		
			[0.019]	[0.047]		
Observations	4192	1622	4112	1498	1430	1081
R-squared	0.04	0.06	0.38	0.38	8.50	8.42
Test stats for $\alpha_1=\alpha_2=\alpha_3=\alpha_4=0$	0.40	1.29	1.86	1.16	(0.075)	(0.077)
(P-value)	(0.809)	(0.229)	(0.135)	(0.341)	15.32	6.94
Test stats for $\beta_1=\beta_2=\beta_3=\beta_4=0$	0.28	1.50	1.12	1.05	(0.004)	(0.139)
(P-value)	(0.890)	(0.218)	(0.360)	(0.391)	1430	1081

Note: Standard errors in brackets are clustered by the state of residence at marriage. Controls included in the regressions but omitted from the table are dummy variables for the year of marriage and for the state of marriage. Test statistics are from F-statistics for columns (1) and (2) and Wald chi2 for columns (3) and (4).

Appendix Table B3: Effects of gender-specific unemployment rates at the time of marriage on the number of children, by age at marriage (Supplement to Table 7)

Sample and dependent variable:	SIPP: the number of children at survey for women who were 35-37 years old at survey		PSID: the number of children at age 35
Unemp. rates at:	Marriage	Age 18-20	Marriage
Female unemployment rate * wife's age at marriage			
20 years old or younger (α_1)	-0.045	-0.061	0.000
	[0.038]	[0.038]	[0.000]
21-23 years old (α_2)	-0.001	-0.012	0.020
	[0.023]	[0.040]	[0.029]
24-27 years old (α_3)	-0.026	-0.026	0.030
	[0.026]	[0.039]	[0.037]
28 years old or older (α_4)	-0.021	-0.033	0.053
	[0.029]	[0.033]	[0.076]
Male-female gap in unemp. rate * wife's age at marriage			
20 years old or younger (β_1)	0.028	0.061	0.000
	[0.029]	[0.043]	[0.000]
21-23 years old (β_2)	-0.069*	-0.068	-0.050
	[0.036]	[0.043]	[0.055]
24-27 years old (β_3)	-0.016	0.000	0.119*
	[0.040]	[0.045]	[0.061]
28 years old or older (β_4)	-0.101**	0.037	-0.135**
	[0.045]	[0.052]	[0.056]
Observations	2,806	2,530	946
Test stats for $\alpha_1=\alpha_2=\alpha_3=\alpha_4=0$	0.84	0.83	0.46
(P-value)	(0.506)	(0.511)	(0.712)
Test stats for $\beta_1=\beta_2=\beta_3=\beta_4=0$	2.93	2.11	3.01
(P-value)	(0.031)	(0.009)	(0.042)

Note: Standard errors in brackets are clustered by the state of residence at marriage/birth. Controls included in the regressions but omitted from the table are dummy variables for the year of marriage/birth, for the state of marriage/birth, and for the age at marriage (21-23, 24-27, 28-35). The number of observations in the second column is fewer than that in Table 7 because women who have not married are excluded.

Appendix Table B4: Effects of the female unemployment rate and the male-female gap in the unemployment rate on the marriage hazard, by state's marriage laws.

	Divorce law favorable for housewives	Divorce law unfavorable for housewives	Common law marriage recognized	Common law marriage not recognized
Female unemployment rate* woman's age				
17-20 years old	0.158***	0.117***	0.104***	0.143***
	[0.019]	[0.020]	[0.015]	[0.016]
21-23 years old	0.061***	0.036***	0.018	0.057***
	[0.016]	[0.013]	[0.013]	[0.012]
24-27 years old	-0.074***	-0.079***	-0.105***	-0.058***
	[0.018]	[0.020]	[0.014]	[0.015]
28 or older	-0.267***	-0.218***	-0.298***	-0.195***
	[0.026]	[0.039]	[0.031]	[0.031]
Male-female gap* woman's age				
17-20 years old	-0.108***	-0.070*	-0.089***	-0.098***
	[0.029]	[0.042]	[0.030]	[0.032]
21-23 years old	-0.025	0.022	-0.027	0.007
	[0.018]	[0.016]	[0.021]	[0.017]
24-27 years old	0.082***	-0.007	0.023	0.052***
	[0.022]	[0.020]	[0.037]	[0.018]
28 or older	0.104**	-0.098**	-0.051	0.001
	[0.050]	[0.040]	[0.095]	[0.043]
Observations	116,353	150,832	69,825	197,360
Persons	17,631	22,852	11,390	29,248

Note: Standard errors in brackets are clustered by the state of residence. The baseline hazard depends on age and is stratified by the year of birth. All columns include controls for state fixed effects.