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Economic inequality and marriage formation

Xue Li
Louisiana State University and Agricultural and Mechanical College

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ABSTRACT

My dissertation primarily investigates the causal impact of economic inequality on marriage formation. I demonstrate how economic inequality among men affects an individual woman's propensity to get married in both the U.S. and China. Based on the framework of Loughran (2002) and Gould and Paserman (2003), I identify the causal impact of male wage inequality on the marriage propensity among women in the U.S. using the 1990 and 2000 Censuses as well as the 2007 American Community Survey. I address the endogeneity and reverse causality problems by applying skill-biased technological shock as an instrument for the wage gap between high and low educated men following the example of Mocan and Unel (2011). I discover that a low educated woman's marriage propensity becomes lower but a high educated woman's marriage propensity becomes higher when there is an increase in the wage ratio between high and low educated men. Additionally, I examine whether in China the income inequality among men affects female marital decision-making by utilizing the China Health and Nutrition Survey (CHNS). I find that a one-standard-deviation increase in the Gini coefficient of male income is associated with an increase in the probability of being "ever married" by 5.8 percentage points for urban women and by 6.9 percentage points for rural women aged 20 to 34 from 1989 to 2009.

CHAPTER 1. INTRODUCTION

Economists care about marriage formation because marriage forms the basis for the basic unit of society, the family. The propensity of young adults to get married is naturally associated with women's fertility choices and children's development as well as people's decisions and outcomes in the labor market. Pioneered by Becker (1973), the causes of marriage formation trends have been widely discussed.

There have been dramatic changes in family formation across the world in the last several decades. Particularly, fewer Americans are getting married today compared with those in the previous 6 to 7 decades (Greenstone and Looney 2012). The percentage of married households among all types of households has dropped from more than 75% in the 1940s to less than 50% in the 2010s based on the statistical results from the U.S. Census Bureau.[1] On the other side of the world, the economic liberalization policies[2] in China have challenged traditional Chinese values and concepts of family during the last 30 years. With the ending of the institution of arranged marriage[3], marriage formation in China is beginning to reflect similar traits as those of western countries.

Not just differences in marriage formation but changes in economic inequality have also caught researchers' attention. There is an ongoing trend of rising economic inequality in recent decades in both the U.S. and China (Yang 1998; Gustafsson and Li 2002; Meng 2004; Autor, Katz and Kearney 2008). Studies have linked the rising earning inequality to factors such as globalization, technological changes and other changes in the labor institutes (Leamer 1996; DiNardo, Fortin and Lemieux 1996; Feenstra and Hanson 1996; Acemoglu 1998; Card and DiNardo 2002). Most of these factors are not in the control of individuals. The rising earning inequality has severe consequences on the society, such as more crime and lower happiness/satisfaction in daily life (Thorbecke and Charumilind 2002; Mocan and Unel 2011). Although people in the labor market with lower earnings seem to marry less (Greenstone and Looney 2012), we are not certain of whether and why there is a causal impact of economic inequality on marriage formation.

This dissertation contributes to existing literature by showing how economic inequality affects marriage formation. I put together two essays to present evidence from the U.S. and China respectively. Both essays approach this problem by exploring the impact of earning inequality among men on the propensity of women to get married.

This approach originates from the literature that links female marital decision-making to male wage inequality in the U.S. (Loughran 2002; Gould and Paserman 2003). They find that women will delay marriage under a higher male wage inequality. This is because a woman usually has a "reservation wage", the minimum acceptable wage of her potential partner during her marital search. The level of the "reservation wage" becomes higher when male wage inequality becomes larger in a U.S. metropolitan area in a year. However, the endogeneity and reverse causality problems in identifying the causal impact have not yet been fully

[1]See *Table HH-1. Households, by Type: 1940 to Present* at www.census.gov/hhes/families/data/households.html

[2]A nation-wide reform and opening up campaign was launched by the Communist Party of China (CPC) in 1978. It is aimed to transit a highly centralized planned economic system to an innovative socialist market economic system, and transmit the closed or semi-closed state to a fully opened state to the outside world.

[3] Since the enactment of 1950 "Marriage Law", arranged marriage has been abandoned in China.

addressed. Neither is there evidence from other parts of the world to support these results (Kuo 2008).

Therefore, in this dissertation, I extend their research in two directions. In the next chapter, I identify the causal impact of male wage inequality on American women's marriage propensity by applying the instrumental variable method. I use skill-biased technological change as the instrument for male wage inequality. Male wage inequality is measured by the wage ratio between high educated (with some college education or higher) men and low educated (with high school degree or lower) men. Based on the data from the 1990 and 2000 U.S. Censuses as well as the 2007 American Community Survey, I find that an increase in the wage gap between high and low educated men decreases the marriage propensity among low educated women and it increases the marriage propensity among high educated women. The same results are found in white or non-white women samples. In general, the findings could explain about 1/5 of the marriage trends in broad educational groups for women in the past 20 years. Two views help explain the mechanism behind the results. One is from the assortative mating by education, which states high educated women usually match with high educated men, while low educated women usually match with low educated men. The other is consistent with Watson and McLanahan (2011). They find that people have a tendency to compare expected household income after marriage with the median household income in the metropolitan area to decide whether to get married within a year.

In the third chapter, I examine the impact of male income inequality on the marriage propensity among urban and rural women in China. I adopt data from the China Health and Nutrition Survey (CHNS). I find that Chinese women aged from 20 to 34 are more likely to be married in an urban city or a rural county[4] with a higher male income inequality. The result is conditional on a woman's personal characteristics including age and education and marriage market features such as the sex ratio between men and women. It is stronger for rural women than for urban women, for elder women than for younger women and for lower educated women than for higher educated women. It is robust under different measurements and specifications. It is also true by using overall income inequality in the city or county. Similar effects are found for men. I further provide two explanations for these results. First of all, when there is higher income inequality among men, especially under polarized income distribution, it is more likely to meet partners with income exceeding the reservation value, which is women's expected standard on male income in marital search. In addition, higher income inequality raises income risk and uncertainty of the society. Hence, with an inadequate social security system, higher income inequality in a city/county makes marriage more attractive to both young women and young men in China, especially in rural areas.

Finally, in the last chapter, I summarize and contrast the findings in both essays and further point out the implication of these results. Before starting the first essay, I have to clarify that in both U.S. and China, marriage formation has never been only attributed to certain financial aspects. The discussion of all the aspects that affect marriage is out of the scope of this study. Instead, this dissertation solely focuses on the impact of earning inequality on the propensity to marry. Not only that, the assumption of the rationality of man may be easily violated in the marriage formation process, as marital decision-making could be highly emotional and involve

[4] "Province, Prefectures, Counties and Townships" is the basic four-level administrative division of China. The "county" or formally "county level division" in China is the third level administrative hierarchy. A "city" in China can be a province-level municipality, a prefecture-level city and a county-level city. In this study, there are samples from cities in urban area and from counties in rural area.

many factors such as “love”, “personality” and “culture” that are hard to quantify. Consequently, the conclusion in this paper is only about one determinant in the marriage formation and does not necessarily apply to any particular person.

CHAPTER 2. SKILL-BIASED TECHNOLOGICAL CHANGE, MALE WAGE INEQUALITY AND FEMALE MARITAL DECISION

2.1 Introduction

This paper investigates whether male wage inequality affects female marriage propensity in the U.S. An increase in the male wage inequality is associated with a decrease in female marriage propensity (Loughran 2002; Gould and Paserman 2003). Theoretically, a woman searches for a husband based on the wage distribution of men conditional on other male characteristics. When male wage inequality increases, due to a higher expected benefit to continue the marital search, the woman forms a higher "reservation wage", which is defined as the minimum acceptable wage level of the potential husband as in the "job search theory". A higher "reservation wage" reduces the probability for the woman to meet a qualified mate, and hence delays marriage.

Current research however cannot identify the causal impact due to the difficulty in addressing endogeneity and reverse causality problems. For instance, some unobserved socioeconomic changes in a metropolitan area such as earning instability and/or income mobility may change both male wage inequality (Gottschalk and Moffitt 1994) and female marriage propensity (Hess 2004). The converse may be true: as the propensity to get married among women changes, the number of married and single men in the metropolitan area also changes. This magnifies male wage inequality, since married men usually earn a wage premium by virtue of being married (Ahituv and Lerman 2007). Previous studies only address these problems by controlling for the metropolitan area fixed effect, year fixed effect and the metropolitan area specific year trends (Loughran 2002; Gould and Paserman 2003; Kuo 2008).

In order to identify the causal impact, I instrument male wage inequality with skill-biased technological change. Skill-biased technological change has compounded the wage gap between high and low educated men since the late 1970s in the U.S. (Card and DiNardo 2002; Autor, Katz, and Kearney 2008). However, general skill-biased technological change in a metropolitan area cannot determine a woman's marriage decision unless through the change of male wage inequality after controlling for her personal characteristics including race, age, education and occupation.

To facilitate the application of this instrument, I measure male wage inequality using the wage gap between high and low educated men[5]. Loughran (2002) and Gould and Paserman (2003) emphasize the role of wage inequality between men within the same educational group rather than the wage gap between high and low educated men in explaining female marriage propensity. Recently, through simulation, Watson and McLanahan (2011) find that couples compare their expected household income with the median household income in their own metropolitan area ("norm") in deciding to get married. Specifically, a lower expected household income compared with the "norm" usually discourages a couple from marriage, and vice versa. A larger wage gap between high and low educated men indicates a relatively lower wage for low educated men in general compared with the median wage of all men, and the opposite holds true for high educated men. Combined with the fact that women usually associate with men in the same broad educational group (Browning, Chiappory, and Weiss 2011), low educated women

[5] In this paper, high educated (skilled) is always defined as with some college or more education and low educated (skilled) otherwise.

are generally less likely to marry conditional on their own earnings, while the opposite should happen for high educated women.

Consistent with this prediction, I find that an increase in the wage gap between high and low educated men decreases the propensity of low educated women to get married, and increases the propensity of high educated women to get married. I use data from the 1990 5% and the 2000 5% U.S. Censuses as well as the 2007 1% American Community Survey (ACS). The result indicates that the roughly 10% increase in the wage ratio between high and low educated men from 1990 to 2007 decreases the rate of low educated women who have been married among all low educated women by at least 4.03 percentage points and increases that for high educated women by at least 2.26 percentage points for women aged 21 to 35 across 212 U.S. metropolitan areas. It accounts for about 22% (4.03%/18%) of the actual decline (18%) in the marriage rate among low educated women. And without the effect of a greater wage gap between high and low educated men, the marriage rate decline among high educated women (12%) would have been almost 19% (2.26%/12%) higher.

Similar to Gould and Paserman (2003), I control for female personal characteristics such as age, race, and education, the sex ratio in specific racial and educational group, and the metropolitan area fixed effects as well as the metropolitan area specific time trends. I construct the instrument following an example from Mocan and Unel (2011), where they instrument low educated men's wage with skill-biased technological change to explain the propensity to conduct crime. I check the robustness with alternative variable measures and by further controlling for the woman's occupation. I also divide sample by race. The results can also apply for white and non-white women separately. After adding the number of prisoners in a state in each year as an additional control, the effects of the male wage gap on female marriage propensity still exist but are generally reduced. Hence, I cannot exclude the possibility that wage gap between high and low educated men alters people's propensity to conduct crime which interacts with women's propensity to get married.

The rest of the chapter is arranged as follows: Section 2.2 review the literature on general determinants of female marriage propensity and specifically the effect of the male wage inequality or the wage gap between high and low educated men on female marriage propensity. In Section 2.3, I explain the usage of skill-biased technological change as an instrument for the wage gap between high and low educated men. Section 2.4 introduces the data and the estimation sample. In Section 2.5, I describe the empirical methodology, while Section 2.6 presents the estimation results, interpretation and robustness checks. Section 2.7 concludes this chapter.

2.2 Literature Review

In this section, I review previous research examining the determinants of female marriage propensity. I focus on the literature linking male wage inequality to female marriage propensity. I also introduce studies that imply a correlation of the wage gap between high and low educated men with female marital decision-making.

2.2.1 Effect of Traditional Factors on Female Marriage Propensity

From an economic perspective, Becker (1973) was the first to study when and why women get married. Under his framework, researchers have been investigating the factors that determine women's marriage decisions. As he observed, women estimate the net benefit of marriage in choosing between singlehood and marriage. Women with different socioeconomic conditions have different preferences toward marriage; most directly, a woman's characteristics, for example: age, race, education and wage, affect her own marriage decision.

To be specific, the propensity to marry increases as a woman gets older. White women are more likely to get married than non-white women. Women may delay their first marriage due to school enrollment, military service, or prolonged cohabitation (Coale 1992; Oppenheimer 1994, 2003; Akerlof, Yellen, and Katz 1996; Goldin and Katz 2002). Women may also have later or fewer marriages as they become more economically independent, where economic independence can be reflected in increased labor force participation, higher education, and improved earnings (Santow and Bracher 1994; Oppenheimer 1997; Okun 2001; Raymo 2003). However, women with higher education and improved financial condition bring more benefit to a marriage union, which may also make it easier for them to get married (Hess 2004; Stevenson and Wolfers 2007).

A woman's marital decision is also affected by marriage market conditions. A marriage market can be viewed as the pool of prime-aged men and women in a particular area in a year. The easier it is to meet a desirable mate in a marriage market, the more likely it is for a woman to get married. The marriage rate of young women increases, given a higher sex ratio, which is calculated by the number of prime age men over women in a metropolitan area in a year (Muhsam 1974; Schoen 1983; Oppenheimer 1988; Angrist 2002). Similarly, a woman is more likely to marry with a higher sex ratio in her specific living, working, or studying environment, more frequent marital search, and the improvement of transportation or telecommunication tools in the local area (Oppenheimer 1988; Xu, Qiang and Wang 2003; Stevenson and Wolfers 2007).

2.2.2 Effect of Male Wage Inequality on Female Marriage Propensity

Besides all of the previously discussed factors, male financial condition also has a noticeable impact on the female marital decision. A number of studies have shown that a lack of financially attractive men reduces the marriage rate among young women (Becker 1981; Wilson 1987; Lichter, LeClere, and McLaughlin 1991; Litcher et al. 1992). Most of these papers use the absolute value of wage or income to measure a man's financial attractiveness. Nevertheless, they imply that the wage comparison among men also matters for a woman's marital decision.

Specifically, recent studies show that male wage inequality affects female marital decisions. They explore the role of male wage inequality during the process of a female marital search. During marital search, a woman searches for a suitable husband among men with different wages conditional on other male characteristics. Generally speaking, women need to spend more time to select a proper mate when income levels among men become more diverse and differentiated in a marriage market (Becker 1974; Keeley 1977). The longer search delays marriage. In addition, as Loughran (2002) points out, when male wage inequality is higher, women also tend to be more discriminating in male wages. Based on the work of Mortensen (1986), he argues that as the distribution of male wage becomes fatter tailed, a woman forms a higher standard on the minimum male wage she can accept in order to get married, in other words, a higher "reservation wage". With a higher reservation wage, the probability to meet someone exceeding this new "reservation wage" becomes smaller if the distribution of the male wage is well shaped with more mass in the center than the tails. Hence, a higher reservation wage normally will lead to later and fewer marriages.

This reservation wage mechanism can be used to explain the empirical results found in the U.S. A higher male wage inequality, especially within a certain educational and racial group, delays marriages for women in the same educational and racial group in a metropolitan area (Loughran 2002; Gould and Paserman 2003; Coughlin and Drewianka 2011). However, in Taiwan, the effect is very small and insignificant (Kuo 2008). These studies usually measure wage inequality with the standard deviation of male wage among men of a certain educational and racial group, or that among all men. Alternative measures such as "Gini coefficient", "90/50" or "50/10" ratio are tested in the robustness checks.

However, these empirical studies have difficulties in claiming a causal relationship, since they are facing unobserved heterogeneity and reverses causality problems (Gould and Paserman 2003). There can be uncontrolled factors that affect both the male wage inequality and female propensity to marry. One example is the earning instability: higher earning instability increases the variance of short-term "transitory" change in earnings, and can account for a higher wage inequality in a certain year (Gottschalk and Moffitt 1994). Higher wage instability may reflect higher job mobility in a metropolitan area. As changes in job positions and earnings are more frequent, marriage may become harder to establish. In this sense, earning instability is also associated with fewer marriages. It is also possible, however, that an earning instability can promote marriage. For instance, women may get married in order to obtain more social supports or share financial risk when there is greater financial instability in society (Hess 2004). As a result, earning instability, for which is hard to control, causes an unobserved heterogeneity problem.

Besides, there can be reverse causality. It is reasonable to doubt that wage inequality among men is affected by female marriage propensity. The effect can be through the "wage premium" for married men (Korenman and Neumar 1991; Chun and Lee 2001; Ahituv and Lerman 2007). As female marriage rates change in a metropolitan area over a period, the numbers of married and single men also change. This also leads to different levels of male wage inequality. Therefore, through "wage premium" effect, marriage propensities among women affect the wage inequality among men.

2.2.3 Effect of Wage Gap between High and Low Educated Men on Female Marriage Propensity

Previous papers mainly address the endogeneity and reverse causality by adding metropolitan fixed effects and metropolitan-specific time trends. In this paper, in addition to controlling for these effects, I use skill-biased technological change as an instrument for male wage inequality. To apply this instrument, I use the wage gap between high and low educated men as the measure for male wage inequality.

Two phenomena together could explain why the wage gap between high and low educated men affects the female marital decision-making. Firstly, the comparison of expected household income after marriage with the median household income of all families within a metropolitan area in a year is important in a couple's decision to marry. If the expected household income is below the median, the couple is less likely to marry and vice versa. In addition, if it is become further above or closer to the median (if not above), the couple are more likely to marry (Watson and McLanahan 2011). Accordingly, one can predict that, conditional on a woman's own earnings, a woman is less (or more) likely to get married with a partner whose wage is further below (or above) the median male wage in a metropolitan area in a year. Therefore, as the wage gap between high and low educated men is magnified, the average income of all low (or high) educated men is further away from the median wage of all men. Therefore, women who match with low educated men become less likely to marry, while women who match with high educated men become more likely to marry.

Secondly, there is assortative mating by education meaning that people usually match with others of similar education. Economists usually explain the economic logic of assortative mating by “transfer utility model” under the framework of Becker’s “bilateral match theory” (Choo and Siow 2006; Seitz, Siow and Choo 2008; Siow 2009). Empirically, educational assortative mating in U.S. has been prevalent for some time (Mare 1991; Pencavel 1998; Lowis and Oppenheimer 2000; Browning et al. 2011).

Therefore, the wage gap between high and low educated men can affect marriage propensity among low and high educated women. We predict that, conditional on the median male wage, low educated women are less likely to get married as wages of low educated men on average are further below the median male wage; whereas high educated women are more likely to get married as wages of high educated men on average are further above the median male wage.

There are still endogeneity and reverse causality problems to be addressed using the wage gap between high and low educated men as a measure for male wage inequality. Generally speaking, unobserved or uncontrolled factors that cause endogeneity for overall male wage inequality can cause similar problems for the wage gap between high and low educated men. To solve these problems, I instrument the wage gap between high and low educated men with an index for skill-biased technological change.

2.3 Instrument

I use a measure of skill-biased technological change as the instrument for the wage gap between high and low educated men. Skill-biased technological change is one of the main drivers of the wage gap between high and low educated workers (Card and DinNardo 2002; Autor et al. 2008). Skill-biased technological change increases the relative demand for high educated workers over low educated workers, and thus leads to a higher wage ratio between high and low educated workers (Autor et al. 2008; Acemoglu and Autor 2010). This explains the persistent increase in the wage gap between high and low educated workers since the late 1970s, despite of a large increase in the relative supply of high educated workers over low educated workers. Mocan and Unel (2011) provide a detailed explanation on the usage of skill-biased technological change as an instrument for wage of low educated workers to explain the propensity for crime. I construct the instrument based mainly on their paper.

An index of skill-biased technological change can be interpreted from a CES production function with inputs of both high skilled and low skilled workers, given the relative supply and the relative wage of high skilled and low skilled workers. Specifically, consider a CES production function (2.1), where high or low educated workers are imperfect substitutes.

$$(2.1)\quad Y = [(A_L L)^{\frac{\sigma-1}{\sigma}} + (A_H H)^{\frac{\sigma-1}{\sigma}}]^{\frac{\sigma}{\sigma-1}}$$

In the function (2.1), H and L represent the efficiency-adjusted labor inputs from high and low skilled workers respectively, while A_H and A_L denote the respective factor-augmenting technology terms for high and low skill labor. The parameter σ represents the elasticity of substitution between high skilled and low skilled workers. A skill-neutral technological change increase A_H and A_L by the same amount which can be reflected by $\sigma = 1$ in the model, whereas a skill-biased technological change will increase A_H more than A_L and hence increase the value of A_H/A_L. As a result, there is $\sigma > 1$ under the assumption of skill-biased technological change (Acemoglu 1998; Autor et al. 2008).

Based on this production function, we can obtain the relationship among relative wage, relative supply and relative demand for high skilled workers compared with low skill workers. Assuming that the labor markets are competitive, low or high skill workers' unit wage equals the value of marginal product of low or high skill workers respectively as shown in equation (2.2) and (2.3) below.

$$(2.2)\ \ \omega_L = \frac{\partial Y}{\partial L} = A_L^{\frac{\sigma-1}{\sigma}}[A_L^{\frac{\sigma-1}{\sigma}} + A_H^{\frac{\sigma-1}{\sigma}}(\frac{H}{L})^{\frac{\sigma-1}{\sigma}}]^{\frac{1}{\sigma-1}}$$

$$(2.3)\ \ \omega_H = \frac{\partial Y}{\partial H} = A_H^{\frac{\sigma-1}{\sigma}}[A_H^{\frac{\sigma-1}{\sigma}} + A_L^{\frac{\sigma-1}{\sigma}}(\frac{H}{L})^{-\frac{\sigma-1}{\sigma}}]^{\frac{1}{\sigma-1}}$$

Then, the wage ratio between high and low skilled (educated) workers or the wage premium for high educated workers can be calculated as

$$(2.4)\ \ \frac{\omega_H}{\omega_L} = \left(\frac{A_H}{A_L}\right)^{\frac{\sigma-1}{\sigma}}(\frac{H}{L})^{-\frac{1}{\sigma}}$$

According to equation (2.4), the change in the wage premium for high educated workers denoted by ω_H/ω_L can be viewed as the result of both the change in relative supply of labor, mainly the value of H/L , and the change in the relative demand of labor, mainly the value of A_H/A_L. The relative supply of labor, H/L, has been rising, which will drive down the wage premium ω_H/ω_L. The increase of ω_H/ω_L over time is mainly due to the rising relative demand, captured by the rise of A_H/A_L. Specifically, skill-biased technological change increases the term A_H/A_L, and drives up the relative demand of high skilled workers.

Accordingly, I use $\ln(A_H/A_L)$ as the measure for skill-biased technological change[6], the instrument. I use the logarithm of the wage ratio between high and low educated male workers, $\ln(w_H/w_L)$, to denote the wage gap between high and low educated men, the key explanatory variable. Implicitly, I assume skill-biased technological change drives up the wage gap between high and low educated male workers in the same way as it does for all workers, male and female workers.

We observe the relative supply and relative wage of high and low skilled workers in a metropolitan area in a year. Hence, the instrument measure can be backed out from equation (2.5). Take the logarithm term for both sides of equation (2.4), we have

$$(2.5)\ \ ln\left(\frac{\omega_H}{\omega_L}\right) = \frac{\sigma-1}{\sigma}ln\left(\frac{A_H}{A_L}\right) - \frac{1}{\sigma}ln(\frac{H}{L})$$

[6] See Mocan and Unel (2011) for more detailed explanation.

Specifically, $\ln(\omega_H/\omega_L)$ is calculated from wage of all full time[7] workers, male and female aged between 17 and 65 in each metropolitan area in a year. H and L are the total efficiency-adjusted hours worked for high and low educated workers respectively in each metropolitan area in a year. I adopt the value of σ used in Unel (2010) and Mocan and Unel (2011), where $\sigma = 1.6$. The instrument variable, $\ln(A_H/A_L)$, is calculated for each metropolitan area in each year.

According to the construction of the instrument, we could believe that skill-biased technological change certainly affects the explanatory variable of interest, the logarithm of wage ratio between high and low educated male workers. But it does not affect the dependent variable, a woman's decision to get married. Consider two women with same age, education and occupation in the same metropolitan area. Skill-biased technological change over time cannot explain why one of them gets married but the other stays single, unless it is because of the change in the wage gap between high and low educated men. Therefore, skill-biased technological change can be a valid instrument for the wage gap between high and low educated men in explaining female marriage propensity, conditional on the woman's age, education and occupation, as well as controlling for metropolitan area fixed effects and metropolitan area specific time trends.

2.4 Data[8]

I use samples of the 1990 and the 2000 Censuses data as well as the 2007 ACS data[9]. The numbers of observations are more than 12.5 million in the 1990 Census, 14.1 million in the 2000 Census and around 3 million in the 2007 ACS. The main variables in use are year, state, metropolitan area, age, sex, marital status, race, education, employment status, weeks worked, hour worked, income wage, occupation, personal weight and migration status.

Previous literature (Loughran 2002; Gould and Paserman 2003) uses samples from the 1980, 1990 and 2000 U.S. Censuses. In my paper, I exclude 1980 Census because of severe problems in the constructed values of the male wage ratio and technological shock ratio, which is probably due to a different weight structure in 1980 Census compared with 1990 and 2000 Census[10]. However, there will be a small sample of metropolitan areas only using the 1990 and 2000 Censuses. I consider adding American Community Survey (ACS) samples, which are comparable with census samples. Among 12 waves of ACS from 2001 to 2012, 4 waves including the 2001 to 2004 ACS do not have proper information in identifying metropolitan areas and 5 waves of ACS from 2008 to 2012 do not have a variable to indicate numbers of "Weeks worked last year". Therefore, the choices narrow down to 3 waves of ACS from 2005 to 2007. However, adding data from three consecutive year to two decennial censuses to calculate variables such as the relative efficiency ratio is problematic and the variations in technological

[7] Workers worked full time, also called full hours full time (FHFT), are those whose weekly hours worked are more than 35 hours and weeks worked are more than 40 weeks,.

[8] See Appendix A for details.

[9] CPS March data is not adopted because that the identification of metropolitan areas in the CPS March data may be inconsistent for the metropolitan area with population less than 500,000, due to the large sampling variability.

[10] 1980 Census has "flat weight" which means each individual has the same personal weight value.

shock of three consecutive years are small. Hence, I only use the 2007 ACS in addition to the 1990 and 2000 Censuses, so that the year gaps are similar among the three waves of data.

I drop partially identified metropolitan areas with more than 30% unidentified observations in any wave of the 1990 5% Census, the 2000 5% Census and the 2007 ACS samples, to solve the partial identification problems.[11] According to this criterion, in total, I drop 30 metropolitan areas, as well as observations without metropolitan identification code in those remaining metropolitan areas in each wave (see Table A.1). There are 224 metropolitan areas (MAs) left in 1990, 256 MAs in 2000 and 256 MAs in 2007. I also delete individuals with personal weight value of 0[12]. After combining three waves, there are 212 metropolitan statistical areas consistently identified. I create a panel with the 212 metropolitan areas which are consistently observed over three waves, 1990, 2000 and 2007.

The sample to estimate with includes over 689,000 women aged between 21 and 35 in the 212 metropolitan areas over three sampled years in the U.S. Table 2.1 presents the sample

Table 2.1 Summary Statistics for Women Aged 21-35 by Year

Variables	1990		2000		2007	
	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.
Age	28.671	4.499	28.840	4.585	27.748	4.354
Years of education	12.856	2.580	12.803	2.803	13.401	2.743
Percentage of married	0.680	0.466	0.622	0.485	0.533	0.499
Proportion of white	0.823	0.382	0.810	0.392	0.826	0.379
Sex ratio (21-40 male/21-35 female)	1.297	0.068	1.387	0.068	1.397	0.082
Sex ratio (low edu)	1.344	0.087	1.577	0.107	1.702	0.162
Sex ratio (high edu)	1.265	0.077	1.233	0.076	1.194	0.107
Sex ratio (white)	1.337	0.065	1.442	0.063	1.454	0.095
Sex ratio (white low edu)	1.372	0.091	1.662	0.117	1.810	0.207
Sex ratio (white high edu)	1.314	0.071	1.282	0.079	1.231	0.113
Sex ratio (black)	1.257	0.816	1.342	1.000	1.301	0.494
Sex ratio (black low edu)	1.391	0.988	1.680	2.976	1.580	1.009
Sex ratio (black high edu)	1.077	0.412	1.125	1.110	1.149	0.777
Technological shock ratio	1.212	0.459	1.156	0.500	1.530	0.548
Wage ratio between high and low educated men	1.523	0.150	1.578	1.170	1.679	0.218
Log of wage ratio between high and low educated men	0.416	0.097	0.450	0.105	0.510	0.126
Observation	254,626		235,823		153,285	

Notes: Means of technological shock ratio, wage gap between high and low educated men and its log value are weighted mean values over 212 metropolitan areas included in the baseline estimation in each year. Other variable means are weighted mean values over individual observations. Data are from the 1990 and 2000 5% U.S. Censuses as well as the 2007 1% American Community Survey.

[11] Partially identified problem happens when PUMA (public use micro-data area) encompassed any territory outside a given metropolitan area, the metropolitan area households located in that PUMA do not receive the relevant code in METAREA.

[12] According to IPUMS, "In 1990-2000, some cases have PERWT values of 0. This is a function of the complex sample design used by the Census Bureau." There are 418 cases in 1990 and 1,758 cases in 2000.

statistics by year. It presents the mean values of the main variables in the baseline regression. The dependent variable is marital status, either "ever married" or not, for each individual woman in each year. The main explanatory variable is the constructed value of the wage gap between high and low educated men in each of the 212 metropolitan areas in a year. It is the logarithm ratio of the average wage among high educated male workers over that among low educated male workers in a metropolitan area in a year, denoted by $\ln(\omega_{MH}/\omega_{ML})_{mt}$. The average wage for high or low educated male workers is calculated from the real weekly wage and salaries, composition-adjusted,[13] for the full time high or low educated male workers aged 17 to 65. The instrumental variable is the constructed value of $\ln(A_H/A_L)_{mt}$, as described in the last section.

Individual level variables include individual marital status, age, years of education and race. The means for individual level variables are the average of the variable values for women aged from 21 to 35 years old in each year weighted by the woman's personal weight in the census or ACS. The metropolitan area level variables include the sex ratio, the sex ratio by education and race, the explanatory variable and the instrumental variable in each of the 212 metropolitan areas. As each woman will reside in a certain metropolitan area, the means of these metropolitan area level variables are the average of the variable values weighted by the corresponding woman's personal weight.

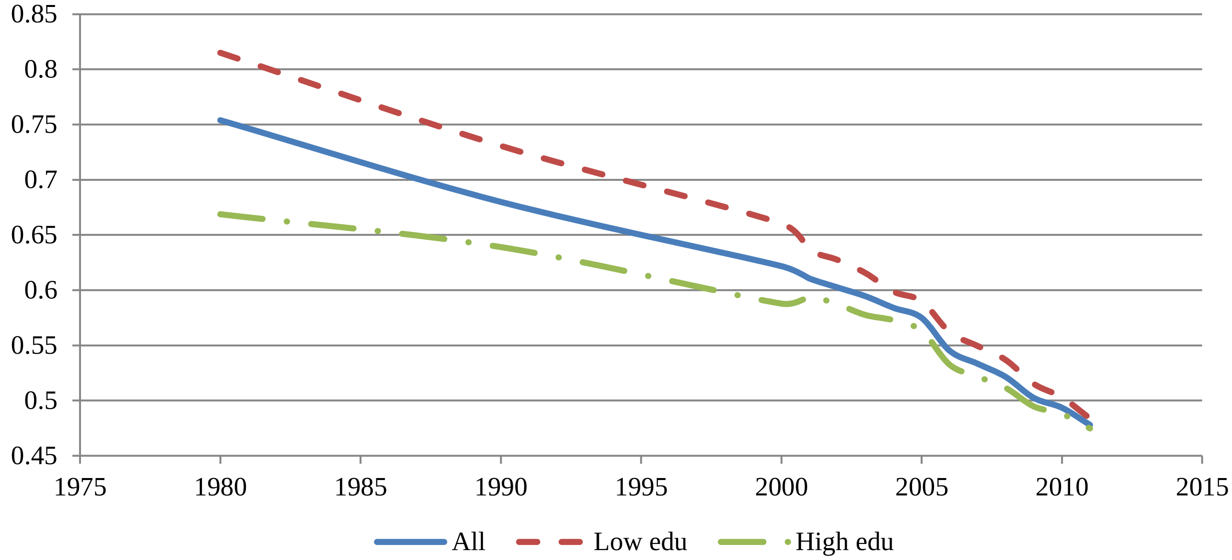
As shown in Table 2.1, the percentage of ever married women of all women aged 21 to 35 is about 68% in 1990. It decreases to about 62% in 2000 and further to about 53% in 2007. The average age of these women is always around 29. The proportion of white women over all women aged 21 to 35 is about 82%. The average education level is about 13 years. The average sex ratio of men aged 21 to 40 [14] over the number of women aged 21 to 35 is 1.3 in 1990 and increases to 1.4 in 2000 and 2007. It increases over time among white people, but not among non-white people. The average sex ratio for low educated people increases from 1.3 in 1990 to 1.6 in 2000 and further 1.7 in 2007. However, it decreases for high educated people from 1.3 in 1990 to 1.2 in 2000 and 2007. This is consistent with the literature that finds high educated women start to outnumber high educated men in recent decades (Browning et al. 2011). The mean value of the wage ratio between high and low educated men increases over the three sampled years, the increase of which is about 10%, from 1.523 in 1990 to 1.670 in 2007. The logarithm values of the wage ratio between high and low educated men are used in the estimation as the main explanatory variable, which on average increases from 0.416 to 0.510 from 1990 to 2007. The mean value of the skill-biased technological change ratio also increases over time.

Figure 2.1 displays some basic trends of marriage rates among women aged 21 to 35 by education. The data are obtained from the 1980 to 2000 5% Censuses and the 2001 to 2010 1% American Community Survey. In general, the proportion of ever married women among all women aged 21 to 35 decreases through 1980 to 2010. Compared to the overall decline in marriage rates for all women, the speed of the decline is relatively slower for high educated women, but faster for low educated women. This is also true for white or non-white women separately. Few studies have explained or even pointed out this phenomenon; however, it agrees with the prediction of the potential links between male wage gap and female marriage propensity discussed in the literature review. In this paper, I suggest that a higher wage gap between high and low educated men increases the marriage propensity for high educated women and decreases

[13] See Appendix A.2 for the construction of composition-adjusted wage.

[14] Consistent with Loughran (2002), as women usually seek men with similar or higher age for future mate, the sex ratio is calculated by the number of men aged 21-40 over the number of women age 21-35.

it for low educated women theoretically. If this effect of wage gap between high and low educated men on high or low educated women is verified empirically, it could provide some explanation for the different speeds of marital decline for high and low educated women.



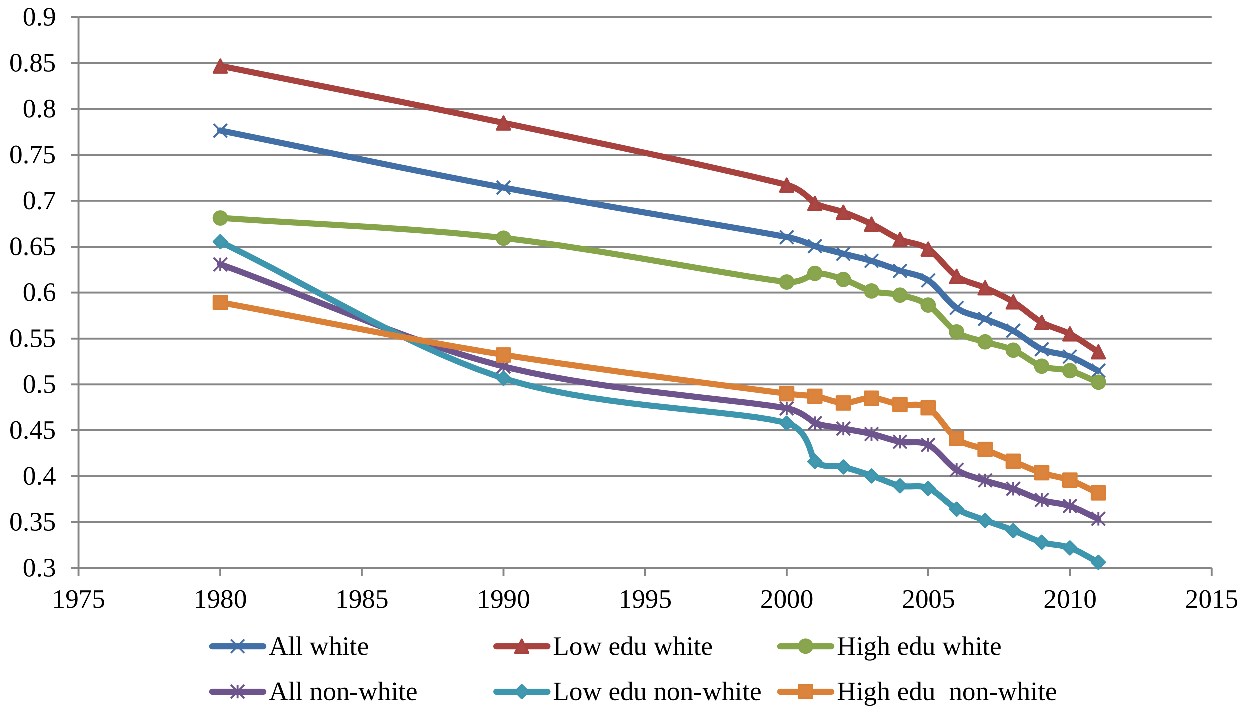


Figure 2.1 Proportion of Women "Ever Been Married" among Women Aged 21 to 35 in the U.S. from 1980 to 2010

Note: Each dot in the graph represents the percentage of ever married women among all women aged 21 to 35 in the U.S. within the defined group. The top graph separates women by educational groups. The bottom graph separates women by educational and racial groups. Data are from the 1980, 1990 and 2000 5% U.S. Censuses and 2001 to 2010 1% American Community Surveys.

2.5 Empirical Specification

The propensity to marry for a woman can be expressed as a function of the wage gap between high and low educated men, woman's personal characteristics, and other related marriage market conditions. Hence, following the literature, I estimate the baseline model (2.6) specified as follows, where I instrument the main explanatory variable, the logarithm of the wage ratio between high and low educated men, with the measure of skill-biased technological change.

(2.6) $Mar_{imt} = a * ln\left(\frac{\omega_{MH}}{\omega_{ML}}\right)_{mt} + b * age_{imt} + c * {age^2}_{imt} + d * edu_{imt} + e * rac_{im} + f * sxr_{mt} + MA_m + T_{mt} + err_{imt}$

Mar_{imt} is a dummy variable to indicate if a woman (i) is ever married in metropolitan area (m) in year (t). This variable has value equal to one if the woman has ever been married including those "currently married, separated, divorced, or widowed" women or equal to zero if the woman has never been married before.

To explain the female marriage propensity, the key variable of interest is the wage gap between high and low educated men in a marriage market, $ln(\omega_{MH}/\omega_{ML})_{mt}$, where I define a marriage market as a metropolitan area (m) in a year (t). According to the mechanism discussed, I expect the sign of the coefficient (a) is negative for low educated women and positive for high educated women.

Since a woman's propensity to get married also depends on her personal characteristics, I add controls for the woman's personal characteristics, including her age, age_{imt}, age squared, ${age^2}_{imt}$, years of education, edu_{imt}, and race, rac_{im}. To be specific, the race dummy, rac_{im}, has value one for whites, and zero for non-whites. The coefficient for race (e) is expected to be positive, because white women are more likely to get married than non-white women. Age and age squared, as well as years of education are all continuous variables. The coefficient of age (b) is expected to be positive, since women become more likely to get married when they get older. The coefficient of education (d) is expected to be negative. Consistent with previous literature, higher educated women usually delay marriage or have lower marriage rates (Goldstein and Kenney 2001).

Since wage is another endogenous variable for propensity to get married, it is not directly controlled for. Instead, years of education is usually used as a proxy for wage (Okun 2001; Goldstein and Kenney, 2001). In this way, we can also keep all women including the non-working women in the sample. Although women's own wage also affects their marital decision, as explained in the literature review the impact is mixed and the results found are usually small and insignificant (Santow and Bracher 1994; Oppenheimer 1994, 1997; Goldstein and Kenney 2001). Nevertheless, one may be still concerned that the instrument, skill-biased technological change, can affect female wage, which will violate the exogenous condition. To further address this concern, in addition to the education control, I add occupation dummies in the robustness checks.

I control for the observable marriage market conditions, approximated by the sex ratio in a metropolitan area in a year. The sex ratio is usually defined as the number of men over the number of women in a specific demographic group. The age range I adopt to calculate the sex ratio is for men 21 to 40 and for women 21 to 35, as women usually prefer to marry men with same or greater age (Choo and Siow 2006). An alternative choice is to use the sex ratio over single men and women. I use the sex ratio for corresponding racial and educational group for sub-samples of women in groups separated by education level and race, which I call "within-group sex ratio". For instance, I regress the marriage propensity for high educated white women on the sex ratio among high educated white men aged 21 to 40 and high educated white women aged 21 to 35. A higher sex ratio indicates that there are more prime age men than women. Hence, I expect the general sex ratio or the within-group sex ratio in a metropolitan area in a year has a positive effect on female marriage propensity.

Furthermore, I control for metropolitan area fixed effect, MA_m, to account for the effect of unobserved metropolitan area conditions that are not changing over time. Finally, I add metropolitan area specific time trends, T_{mt}, to control for the effect of unobserved factors changing over the linear time trend in each metropolitan area. As a result, I only employ the variation of male wage gap in the same metropolitan area deviated from its linear time trend in each year to identify its causal impact on marriage propensity of women in that metropolitan area in each year.

2.6 Results

2.6.1 Baseline IV Estimation

Table 2.2 presents the baseline IV results of model (2.6) under nine specifications, where the male wage gap is instrumented by skill-biased technological change. The coefficients of the instrument in first-stage regressions are positive and significant for all nine specifications. A larger skill-biased technological shock significantly increases the wage gap between high and low educated men in a metropolitan area in a year. The F statistics in the first-stage regression of the instrument is about 17. Hence, the instrument is strong. Standard errors are adjusted for clusters defined by women's age, metropolitan area and year due to the grouped structure of the error terms.

Column (1) shows the effect of the wage gap between high and low educated men on the marriage propensity for a woman aged 21 to 35. I then separate samples into two subsamples by education. A change in wage gap (the logarithm of the wage ratio) between high and low educated men does not have significant impact on women's propensity to get married in general. Column (2) presents the result for low educated women, while column (3) presents that for high educated women. After separating women by education, an increase in the wage gap between high and low educated men decreases a low educated woman's propensity to get married, but increases a high educated women's propensity to get married. Further, I separate the female sample by race. Column (4) shows that a change in wage gap (the logarithm of the wage ratio) between high and low educated men does not have significant effect on a white woman's propensity to get married. But after separating high and low educated white women, column (5) and (6) indicate that an increase in the male wage gap decreases a low educated white woman's propensity to get married, but increases a high educated white woman's propensity to get married. Similarly, column (7), (8) and (9) indicate that an increase in the male wage gap does not have significant effect for a non-white woman when pooling low and high non-white women together, but have significant effects when separating the samples for low and high educated women. It decreases a low educated non-white woman's propensity to get married, but increases that for a high educated non-white woman.

Table 2.2 IV Result: Effect of the Wage Gap between High and Low Educated Men on Female Marriage Propensity

Dependent variable: Marital status (=1 if ever married, =0 if never married) for a woman aged between 21-35 in 1990, 2000 and 2007

	Woman (21-35)			White Woman (21-35)			Non-white Woman (21-35)		
	(1) Pool sample	(2) Low educated	(3) High educated	(4) White woman	(5) Low educated	(6) High educated	(7) Non-white	(8) Low educated	(9) High educated
Log of wage ratio between high and low educated men	-0.071	-0.423***	0.237*	-0.084	-0.432***	0.287**	0.077	-0.738***	0.481***
	[0.096]	[0.117]	[0.126]	[0.103]	[0.138]	[0.142]	[0.122]	[0.170]	[0.174]
Age	0.164***	0.150***	0.169***	0.197***	0.188***	0.202***	0.056***	0.044***	0.061***
	[0.005]	[0.005]	[0.006]	[0.005]	[0.006]	[0.007]	[0.006]	[0.008]	[0.008]
Age^2	-0.002***	-0.002***	-0.002***	-0.003***	-0.003***	-0.003***	-0.000**	0	-0.000*
	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]
Years of education	-0.005***	-0.002***	-0.004***	-0.010***	-0.001	-0.012***	0.011***	-0.007***	0.020***
	[0.001]	[0.001]	[0.001]	[0.001]	[0.001]	[0.001]	[0.001]	[0.001]	[0.002]
Race dummy for white	0.167***	0.240***	0.109***						
	[0.004]	[0.005]	[0.004]						
Within group sex ratio	-0.012	0.050***	0.100**	-0.004	0.037**	0.070*	0.080***	0.031	0.119***
	[0.036]	[0.015]	[0.041]	[0.035]	[0.015]	[0.042]	[0.027]	[0.021]	[0.034]
Constant	-2.399***	-2.081***	-2.915***	-2.684***	-2.401***	-3.141***	-1.265***	-0.268	-1.764***
	[0.111]	[0.109]	[0.114]	[0.114]	[0.123]	[0.116]	[0.256]	[0.321]	[0.195]
First stage: Coefficient of technological shock	0.282***	0.219***	0.243***	0.290***	0.215***	0.232***	0.305***	0.280***	0.312***
	[0.009]	[0.010]	[0.008]	[0.009]	[0.010]	[0.008]	[0.009]	[0.009]	[0.010]
Observations	689,034	310,467	378,567	541,524	241,152	300,372	147,510	69,315	78,195
R-squared	0.249	0.224	0.285	0.259	0.203	0.308	0.178	0.149	0.216
Number of clusters	9,539	9,434	9,483	9,536	9,383	9,462	7,647	6,584	6,408

Note: ***p<0.01, **p<0.05, *p<0.1. Standard errors are adjusted by age-metropolitan-year clusters. Control for metropolitan fixed effect with 212 metropolitan dummies and corresponding time trend in each metropolitan area. Wages are composition-adjusted.

The coefficient values of this main explanatory variable display how much the expected mean values of the marriage propensity for a woman (the expected mean probability for her to be ever married) will change when the logarithm value of the wage ratio between high and low educated men increases by 1. To be more informative, the products of ln(101/100) (≈0.01) and the coefficient values are the value changes in the expected means of the probability to be ever married due to a 1% increase in the wage ratio between high and low educated men.

To be specific, based on the IV results using the sample of low educated women aged between 21 and 35, a 1% increase in the wage ratio between high and low educated men decreases a low educated woman's probability to ever be married by about 0.423 (0.423*ln(1.01)) percentage point. However, according to the results from the sample of high educated women aged between 21 and 35, a 1% increase in the wage ratio between high and low educated men causes a high educated woman's probability to ever be married to increase by about 0.237 (0.237*ln(1.01)) percentage point.

Similar effects are found for white and non-white women. For high or low educated white women aged between 21 to 35, a 1% increase in the wage ratio between high and low educated men causes about 0.432 (0.432*ln(1.01)) percentage point decrease in a low educated white woman's probability to marry, but about 0.287 (0.287*ln(1.01)) percentage point increase in a high educated white woman's probability to be ever married. Finally, IV results indicate a 1% increase in the wage ratio between high and low educated men decreases the probability of a low educated non-white woman to marry by 0.738 (0.738*ln(1.01)) percentage point and it increases the probability of a high educated non-white woman to marry by 0.481 (0.481*ln(1.01)) percentage point.

In fact, the wage ratio between high and low educated men increases, on average, 10% from 1990 to 2007 according to Table 2.1. I then obtain the change in the expected mean of a woman's marriage propensity (probability to be ever married) due to a 10% increase in the wage ratio between high and low educated men by calculating the product of the coefficient values of the main explanatory variable and ln(1.1). [15]

Results are shown in Table 2.3 row 1. A 10% increase in the wage ratio between high and low educated men decreases the probability to be ever married by 4.03 percentage points for a low educated woman, and also decreases that by 4.12 and 7.03 percentage points for a low educated white woman or a low educated non-white woman respectively. A 10% increase in the wage ratio between high and low educated men increases the probability to be ever married for a high educated woman by 2.26 percentage points, and also increases that by 2.74 and 4.58 percentage points for a high educated white woman or a high educated non-white woman respectively.

The individual marriage propensity change can be observed by the marriage rate change for women aged between 21 and 35 at an aggregate level. Refer to Figure 2.1, the marriage rate for low educated women declines about 18% from 73% in 1990 to 55% in 2007. The 10% increase in the wage ratio between high and low educated men reduces a low educated woman's propensity to get married by 4.03 percentage points. At an aggregate level, this can reflect a decrease of 4.03 percentage points in the marriage rates among 21 to 35 years old low educated

[15] For simplicity, let equation (2.6) be expressed as $Mar_1 = aln(male\ wage\ ratio) + others$. When male wage ratio increases by 10% and others remain unchanged, there is $Mar_2 = aln(male\ wage\ ratio * 1.1) + others = aln(male\ wage\ ratio) + others + a * \ln(1.1)$. So, we can get $Mar_2 - Mar_1 = a * \ln(1.1)$. Hence, when male wage ratio increases by 10%, the marriage propensity for a woman would increase by the amount of a*ln(1.1)*100 percent, which is the coefficient for ln(male wage ratio) in the estimation of (2.6) times ln(1.1).

women. Therefore, the effect of male wage gap on female marriage propensity explains about 22% (4.03/18) of the marriage rates decline. Marriage rate for high educated women however decreases by about 12% from 64% in 1990 to 52% in 2007. The 10% increase in the male wage ratio between high and low educated men increases a high educated woman's propensity to get married by 2.26 percentage points. At an aggregate level, it can also lead to an increase of 2.26 percentage points in the marriage rate among 21 to 35 years old low educated women. Therefore, based on the actual decline of 12%, the reduction is about 19% (2.26/12). If there is no such effect, the decline should have been 14.26%, which I call the "counterfactual decline". The effect reduces about 16% (2.26/14.26) of this counterfactual decline of marriage rates among high educated women.

Table 2.3 Change in Expected Mean Female Marriage Propensity Due to A 10% Increase in the Wage Ratio between High and Low Educated Men (Unit: Percentage Point)

10% increase in wage ratio between high and low educated men	Woman (21-35)		White Woman (21-35)		Non-white Woman (21-35)	
	Low educated	High educated	Low educated	High educated	Low educated	High educated
Composition-adjusted (Baseline)	-4.03***	2.26*	-4.12***	2.74**	-7.03***	4.58***
Composition-adjusted race-specific			-3.31*	3.92**	-8.23**	15.5**
Efficiency-adjusted	-7.39***	4.46*	-7.14***	5.92*	-15.7***	8.99***
Efficiency-adjusted race-specific			-7.42***	5.15*	-36.95**	-

Note: ***p<0.01, **p<0.05, *p<0.1. Value presented are the percentage points change of a woman's marriage propensity due to a 10% increase in the wage ratio between high and low educated men, calculated by corresponding estimation coefficients times ln(1.1).

Accordingly, one can calculate how much the marriage rate change among white or non-white women can be explained by the effect of male wage gap on female marriage propensity. From 1990 to 2007, the marriage rates decline is about 18% for low educated white women and 11% for high educated white women. The effect of a 10% increase in the wage ratio between high and low educated men increases over the same time explains about 23% (4.12/18) of the marriage rate decline among low educated white women and reduces about 20% (2.74/13.74) of the counterfactual decline (11+2.74) among high educated white women. For non-white women, the marriage rates decline is about 15% and 10% for low and high educated non-white women respectively. Therefore, the 10% increase in the wage ratio between high and low educated men explains about 47% (7.03/15) of the marriage rate decline of low educated non-white women, and reduces about 31% (4.58/14.58) of the counterfactual marriage rate decline. In general, the effects are more influential among non-white women.

The coefficients for the control variables are generally significant with expected signs. Marriage propensity is higher for an older woman, and higher for whites; it is also higher for a woman in the metropolitan area with higher sex ratio within her educational and racial group. Marriage propensity is lower, however, for a woman with more years of education.

2.6.2 Alternative Measures and Subsamples

I use the "race-specific male wage gap" as main explanatory variable to check the robustness of the baseline results, as women mostly marry men within the same racial group (Seitz, 2009). Specifically, I use the logarithm of the wage ratio between high and low educated white men and the logarithm of the wage ratio between high and low educated non-white men respectively as the main explanatory variable for separate samples of white and non-white women. I apply the same instrument as in the baseline estimation.

Table 2.4 presents the IV results. The coefficients of the instrument in first-stage regressions are always positive and significant with F values bigger than 10. Second stage results are generally consistent with the baseline results. The race-specific male wage gap significantly reduces the marriage propensity for low educated white women and non-white women, but significantly increases that for high educated white women and non-white women. Specifically, a 1% increase in the wage ratio between high and low educated white men decreases the expected mean marriage propensity for a low educated white woman by 0.347 (0.347*ln(1.01)) percentage point and it increases that for a high educated white woman by 0.411 (0.411*ln(1.01)) percentage point. For or non-whites, a 1% increase in the wage ratio between high and low educated non-white men decreases the expected average marriage propensity for a low educated non-white woman by 0.863 (0.863*ln(1.01)) percentage point and it increases that for a high educated non-white woman by 1.627 (1.627*ln(1.01)) percentage points.

As presented in Table 2.3 row 2, a 10% increase in the wage ratio between high and low educated white men decreases the average expected marriage propensity for a low educated white woman aged between 21 and 35 by 3.31 percentage points, but increases that for a high educated white woman aged between 21 and 35 by 3.92 percentage points. A 10% increase in the wage ratio between high and low educated non-white men decreases a low educated non-white woman's average expected marriage propensity by 8.23 percentage points, but increases a high educated non-white woman's average expected marriage propensity by 15.5 percentage points.

I also check the robustness of the baseline results with the alternative methods to adjust wage. We first separate workers into different efficiency groups by race, gender, education and experience. The relative wage ratio of each efficiency group compared with one base efficiency group is the relative efficiency factor. Average wage of each group weighted by the relative efficiency factor is called efficiency-adjusted wage. The average wage of each group weighted further by the compositional weight of labor inputs in each group in addition to the efficiency adjustment is called composition-adjusted wage. [16] Previous results are using composition-adjusted wage. Table 2.3 row 3 and 4 present results using efficiency-adjusted wage. Results obtained by efficiency-adjusted wage have the same signs but bigger magnitudes compared with the results obtained by composition-adjusted wage, which is consistent for whites and non-whites separately. For example, 10% increase in the efficiency-adjusted wage ratio between high and low educated men decreases a low educated woman's marriage propensity by 7.39 percentage points, which explains about 41% (7.39/18) of the aggregate marriage rates decline among low educated women. A 10% increase in the efficiency-adjusted wage ratio between high and low educated men, however, increases a high educated woman's marriage propensity

[16] See Appendix A for details.

Table 2.4 IV Result: Effect of the Race-Specific Wage Gap between High and Low Educated Men on Female Marriage Propensity

Dependent variable: Marital status (=1 if ever married, =0 if never married) for woman aged 21-35 in 1990, 2000 and 2007

	White Woman (21-35)		Non-white Woman (21-35)	
	Low educated	High educated	Low educated	High educated
Logarithm of race-specific wage ratio between high and low educated men	-0.347*	0.411**	-0.863**	1.627**
	[0.190]	[0.186]	[0.417]	[0.808]
Age	0.187***	0.201***	0.046***	0.064***
	[0.006]	[0.007]	[0.010]	[0.010]
Age^2	-0.003***	-0.003***	0	-0.000**
	[0.000]	[0.000]	[0.000]	[0.000]
Years of education	-0.001	-0.012***	-0.009***	0.023***
	[0.001]	[0.001]	[0.002]	[0.002]
Within group sex ratio	0.045**	0.053	0.289	-2.322**
	[0.019]	[0.048]	[0.216]	[0.997]
Constant	-2.441***	-3.152***	-0.803	2.009
	[0.143]	[0.117]	[0.549]	[1.494]
First stage: Coefficient of technological shock	0.212***	0.225***	0.064***	1.558***
	[0.010]	[0.008]	[0.000]	[0.000]
Observations	236,052	296,461	44,217	51,682
R-squared	0.204	0.308	0.147	0.213
Number of clusters	9,296	9,373	4,310	4,233

Note: ***p<0.01, **p<0.05, *p<0.1. Standard errors are adjusted by age-metropolitan-year clusters. Key explanatory variable is the logarithm of the wage ratio between high and low educated white men and the logarithm of wage ratio between high and low educated non-white men respectively when the dependent variable is for white or non-white women. Control for metropolitan fixed effect with 210 metropolitan dummies and corresponding metropolitan time trends. Wages are composition-adjusted.

by 4.46 percentage points, which reduces about 27% (4.46/16.46) of the counterfactual decline of marriage rates if there is no such increase effect.

In Table 2.5, I show regression results using efficiency-adjusted wage for samples of all women and white women aged from 21 to 35. Given the same specification, I further check the robustness using only black women. Results for low educated black women are similar to that for low educated non-white women. It becomes insignificant for high educated black women, probably because that there are few samples of high educated black women aged 21 to 35 in some metropolitan areas in some year.

Besides, I check results using alternative measures of the dependent and some control variables. For instance, I delete separated, divorced and widowed women and restrict the sample to include only currently married and single women, but it does not change the basic conclusion. Using women aged 21 to 30 instead of 21 to 35 does not alter the major results, either. Similarly, there are no major changes in the results by using sex ratio with the alternative age range of men and women or the sex ratio among single men and women. .

Table 2.5 IV Result: Effect of the Efficiency-Adjusted Wage Gap between High and Low Educated Men on White and Black Women's Marriage Propensity

Dependent variable: Married now (=1 if ever married, =0 if never married) for women aged 21-35 in 1990, 2000 and 2007

	Woman (21-35)			White Woman (21-35)			Black Woman (21-35)		
	Pool sample	Low educated	High educated	White woman	Low educated	High educated	Black woman	Low educated	High educated
Logarithm of wage ratio between high and low educated men	-0.410**	-0.743***	0.415	-0.183	-0.749***	0.621*	-0.797***	-1.586***	-0.526
	[0.209]	[0.213]	[0.256]	[0.196]	[0.229]	[0.323]	[0.250]	[0.400]	[0.359]
Age	0.163***	0.149***	0.173***	0.197***	0.188***	0.201***	0.016**	0.017*	0.025**
	[0.005]	[0.005]	[0.007]	[0.005]	[0.006]	[0.007]	[0.007]	[0.009]	[0.010]
Age^2	-0.002***	-0.002***	-0.002***	-0.003***	-0.003***	-0.003***	0.000**	0	0
	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]
Years of education	-0.008***	-0.001	-0.010***	-0.010***	-0.001	-0.012***	0.008***	0.004**	0.002
	[0.001]	[0.001]	[0.001]	[0.001]	[0.001]	[0.001]	[0.001]	[0.002]	[0.002]
Race dummy for white	0.252***	0.307***	0.205***						
	[0.005]	[0.005]	[0.005]						
Within group sex ratio	-0.005	0.062***	0.073*	0.006	0.051***	0.078*	0.053**	0.006	0.064**
	[0.034]	[0.013]	[0.038]	[0.031]	[0.013]	[0.040]	[0.025]	[0.024]	[0.027]
Constant	-2.513***	-2.411***	-2.788***	-2.754***	-2.680***	-2.963***	-0.585*	-0.437	-0.346**
	[0.097]	[0.096]	[0.130]	[0.103]	[0.104]	[0.134]	[0.333]	[0.395]	[0.168]
First stage: Coefficient of technological shock	0.060***	0.076***	0.058***	0.068***	0.079***	0.050***	0.059***	0.054***	0.069***
	[0.003]	[0.004]	[0.004]	[0.003]	[0.004]	[0.004]	[0.003]	[0.004]	[0.003]
Observations	643,734	295,500	348,234	541,524	241,152	300,372	102,210	54,348	47,862
R-squared	0.263	0.239	0.296	0.259	0.203	0.308	0.129	0.109	0.162
Number of clusters	9,539	9,424	9,478	9,536	9,383	9,462	6,530	5,658	5,341

Note: ***p<0.01, **p<0.05, *p<0.1. Standard errors are adjusted by age-metropolitan-year clusters. Control for metropolitan fixed effect with 212 metropolitan dummies and corresponding time trend in each metropolitan area. Wages are efficiency-adjusted.

2.6.3 Additional Controls

In this part, I first add additional occupational control to further examine the causal relationship. As a valid instrument, skill-biased technological change should not be correlated with other uncontrolled determinants of female marriage propensity. Skill-biased technological change can affect female wage as well as the wage gap between high and low educated women, which are not directly controlled for in the estimation. Previous literature shows that the impact of female wage on their marriage propensity is mixed or insignificant, and I already control for years of education. Nevertheless, one might still argue that the wage gap between high and low educated women has similar impacts on their marriage propensity as that of male wage gap, since women would have expected household income closer to or further away from the median household income ("norm") in a metropolitan area within a year as the female wage gap changes. In order to address this concern, I add an occupation dummy to the estimation in Table 2.6 to further control for each woman's own wage. If the results for male wage gap also reflect the effect of female wage gap, the coefficient should be insignificant and relatively smaller after adding additional wage control. As shown in Table 2.6, the coefficients for male wage gap are still significant with the same signs as in Table 2.5 and have even larger magnitudes. This indicates the effects found on female marriage propensity are not from the female wage inequality or the woman's own wage but from the male wage inequality.

Previously, the effect of wage gap between high and low educated men on female marriage propensity is explained to be the result of both assortative mating and the women's tendency to compare relative expected household income with the median household income in her metropolitan area. Alternatively, the mechanism can be through the effect of male wage gap on crime in that metropolitan area in a year. This is because a higher male wage gap may lead to a higher propensity to conduct crime among low educated men (Mocan and Unel 2011). A higher crime propensity for low educated men can cause more incarceration among low educated men and also the deterioration of their socioeconomic conditions, such as physical and mental health, social and economic stability and neighborhood conditions. Those factors can decrease low educated women's propensity to marry.

Table 2.7 presents the IV results after further controlling for the number of prisoners in a state within a year, the data for which are available on the website of U.S. Department of Justice. The signs of the coefficients for main explanatory variable are still consistent. But the magnitudes are generally smaller compared with previous results and some are not significant. So it is possible that part of the effects of male wage gap on female marriage propensity reflects the effect of crime on marriage formation.

2.7 Conclusion

This paper enriches literature in investigating the effect of male wage inequality on female marriage propensity. The endogeneity and reverse causality problems of wage inequality are addressed by using an instrument, the skill-biased technological change, for male wage inequality. To apply the instrument, the wage gap between high and low educated men is used as the measure for male wage inequality. To be specific, this paper identifies the causal impact of the wage gap between high and low educated men on female marriage propensity.

Table 2.6 IV Result: Effect of the Wage Gap between High and Low Educated Men on White and Black Women's Marriage Propensity Controlling for Occupation

Dependent variable: Married now (=1 if ever married, =0 if never married) for women aged 21-35 in 1990, 2000 and 2007

	Woman (21-35)			White Woman (21-35)			Black Woman (21-35)		
	Pool sample	Low educated	High educated	White woman	Low educated	High educated	Black woman	Low educated	High educated
Logarithm of wage ratio between high and low educated men	-0.394*	-1.641***	1.173***	-0.173	-1.538***	1.507***	-0.750***	-1.508***	0.023
	[0.206]	[0.204]	[0.293]	[0.194]	[0.228]	[0.330]	[0.249]	[0.304]	[0.347]
Age	0.162***	0.148***	0.175***	0.195***	0.186***	0.204***	0.015**	0.016*	0.025**
	[0.005]	[0.005]	[0.006]	[0.005]	[0.006]	[0.006]	[0.007]	[0.009]	[0.010]
Age^2	-0.002***	-0.002***	-0.002***	-0.003***	-0.003***	-0.003***	0.000**	0	0
	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]
Years of education	-0.007***	0.001	-0.012***	-0.008***	0.002**	-0.014***	0.005***	0.002	-0.002
	[0.001]	[0.001]	[0.001]	[0.001]	[0.001]	[0.001]	[0.001]	[0.002]	[0.002]
Race dummy for white	0.250***	0.306***	0.204***						
	[0.005]	[0.005]	[0.005]						
Within group sex ratio	-0.001	0.02	0.131***	0.012	0.054**	0.109***	0.053**	-0.014	0.027
	[0.034]	[0.022]	[0.048]	[0.031]	[0.023]	[0.042]	[0.025]	[0.015]	[0.027]
Managerial and professional	-0.078***	-0.062***	-0.104***	-0.099***	-0.087***	-0.112***	0.034***	0.052***	-0.026*
	[0.004]	[0.006]	[0.005]	[0.004]	[0.007]	[0.006]	[0.008]	[0.012]	[0.014]
Technical, sales, and administrative	-0.090***	-0.074***	-0.127***	-0.109***	-0.098***	-0.134***	0.009	0.032***	-0.054***
	[0.004]	[0.004]	[0.005]	[0.004]	[0.005]	[0.005]	[0.007]	[0.008]	[0.013]
Service	-0.117***	-0.097***	-0.154***	-0.130***	-0.115***	-0.158***	-0.020***	-0.003	-0.069***
	[0.004]	[0.005]	[0.006]	[0.004]	[0.005]	[0.007]	[0.008]	[0.008]	[0.015]
Farming, forestry, and fishing	-0.097***	-0.047***	-0.195***	-0.116***	-0.060***	-0.206***	-0.004	-0.013	0.017
	[0.011]	[0.014]	[0.018]	[0.011]	[0.014]	[0.018]	[0.045]	[0.046]	[0.104]
Precision production, craft, and repairers	-0.103***	-0.078***	-0.148***	-0.125***	-0.109***	-0.153***	0.024	0.074***	-0.066**
	[0.008]	[0.011]	[0.011]	[0.009]	[0.012]	[0.012]	[0.019]	[0.025]	[0.030]
Operatives and laborers	-0.104***	-0.088***	-0.155***	-0.121***	-0.112***	-0.156***	-0.003	0.027**	-0.072***
	[0.005]	[0.006]	[0.010]	[0.006]	[0.006]	[0.011]	[0.010]	[0.012]	[0.021]
Observations	643,734	295,500	348,234	541,524	241,152	300,372	102,210	54,348	47,862
R-squared	0.267	0.241	0.3	0.265	0.208	0.311	0.13	0.112	0.164

Note: ***p<0.01, **p<0.05, *p<0.1. Standard errors are adjusted by age-metropolitan-year clusters. The left out category of occupation dummies is the "non-occupational responses", mostly including women without work. First-stage results are significant with F values bigger than 10. Wages are efficiency-adjusted.

Table 2.7 IV Result: Effect of the Wage Gap between High and Low Educated Men on Female Marriage Propensity Controlling for Number of Prisoners in State

Dependent variable: Married now (=1 if ever married, =0 if never married) for women aged 21-35 in 1990, 2000 and 2007

	Woman 21-35		White Woman 21-35		Non-white Woman 21-35	
	Low educated	High educated	Low educated	High educated	Low educated	High educated
Baseline: Main explanatory variable is log of composition-adjusted wage ratio between high and low educated men.						
Logarithm of wage ratio between	-0.211*	0.172	-0.242*	0.233*	-0.456**	0.201
high and low educated men	[0.120]	[0.127]	[0.138]	[0.140]	[0.191]	[0.201]
Number of prisoners	0.007***	-0.007***	0.005***	-0.005**	0.011***	-0.014***
(by state-year)	[0.002]	[0.002]	[0.002]	[0.002]	[0.004]	[0.005]
Main explanatory variable is log of efficiency-adjusted wage ratio between high and low educated men.						
Logarithm of wage ratio between	-0.381*	0.322	-0.416*	0.486	-0.970**	0.368
high and low educated men	[0.203]	[0.253]	[0.221]	[0.316]	[0.414]	[0.371]
Number of prisoners	0.007***	-0.007***	0.005***	-0.005**	0.012***	-0.014***
(by state-year)	[0.002]	[0.002]	[0.002]	[0.002]	[0.004]	[0.004]
Main explanatory variable is log of efficiency-adjusted wage ratio between high and low educated men. Non-white women sample only includes black women.						
Logarithm of wage ratio between	-0.408*	0.366			-1.128***	-0.345
high and low educated men	[0.208]	[0.253]			[0.434]	[0.383]
Number of prisoners	0.006***	-0.003*			0.008**	0.004
(by state-year)	[0.002]	[0.002]			[0.004]	[0.004]
Main explanatory variable is log of composition-adjusted race-specific male wage ratio.						
Logarithm of race-specific wage ratio			-0.174	0.338*	-0.534**	-0.105
between high and low educated men			[0.198]	[0.181]	[0.211]	[0.225]
Number of prisoners			0.005***	-0.004**	0.011***	-0.015***
(by state-year)			[0.002]	[0.002]	[0.004]	[0.005]
Main explanatory variable is efficiency-adjusted race-specific male wage ratio.						
Logarithm of race-specific wage ratio			-0.426*	0.496	-0.907**	0.355
between high and low educated men			[0.226]	[0.321]	[0.385]	[0.362]
Number of prisoners			0.005***	-0.005**	0.012***	-0.014***
(by state-year)			[0.002]	[0.002]	[0.004]	[0.004]

Note: ***p<0.01, **p<0.05, *p<0.1. Standard errors are adjusted by age-metropolitan-year clusters. First stage results are all significant with F values bigger than 10. Data for number of prisoners in each state each year are from the website of U.S. Department of Justice. Only the coefficients of IV results for male wage gap and the additional crime control are shown.

The instrumental variable method is applied to a standard model in testing relationship between male wage inequality and female marriage propensity. The model controls for women's personal characteristics such as age and education as well as the marriage market condition approximated by the sex ratio. It also accounts for the metropolitan area fixed effects and the metropolitan area specific time trends.

This paper finds that an increase in the wage gap between high and low educated men in a metropolitan area in a year results in lower marriage propensity for low educated women but a higher marriage propensity for high educated women aged from 21 to 35. The reduction in marriage propensity for a low educated woman is usually higher than the increase in marriage propensity for a high educated woman, in response to the same change in the male wage gap. Basically, a 10% increase in wage gap between high and low educated men decreases marriage propensity for a low educated woman by at least 4 percentage points and increases that for a high educated woman by at least 2 percentage points.

Similar results also persist among white or non-white women separately. For white women, the magnitude is closer to the general results, though for non-white women or only black women, the magnitude is greater than that for white women. Hence, the effect of the male wage gap on the propensity to marry seems to be more influential among non-white women.

The results are robust to alternative sample specifications and variable measurements. For instance, the effect of the race-specific male wage gap on the propensity to marry among whites or non-whites respectively are similar to the effect of the general male wage gap. Results using efficiency-adjusted wage lead to the same signs but larger magnitudes compared with the results using composition-adjusted wage.

The impacts are also economically significant. At the aggregate level, the change in marriage propensity transfers to the change in the marriage rates among high or low educated women aged from 21 to 35. These findings help explain the relative slower decline of high educated women's marriage rates and the relative faster decline of low educated women's marriage rates compared with the general marriage decline trends. For instance, the baseline results explain more than 22% of the overall marriage rates decline among low educated women. Meanwhile, it reduces more than 16% of the counterfactual decline of the marriage rates among high educated women. With efficiency-adjusted method, the figure can go up to 41% for low educated women and 27% for high educated women.

The causal effect of male wage gap on female marriage propensity is channeled by both the comparison of expected household income with the median household income in the metropolitan area and the phenomenon of educational assortative mating. They are not due to the change in female wage or wage gap between high and low educated women. However, the impact of the wage gap between high and low educated men on female marriage propensity may channel through the impacts of male wage gap on crime.

The paper also has important practical implications. It presents another social condition that can be changed by a higher wage inequality. Rising wage inequality among men, potentially driven by the skill-biased technological change in recent decades, has not only led to social changes such as a higher crime propensity among the poor, but also affects the female marriage propensity differently among high and low educated women. It demonstrates how inequality in economic conditions causes an inequality in the marriage formation. Accordingly, as male wage inequality continue to increase, marriage is becoming a privilege for richer and higher educated people, and an impossible dream for poorer and less educated people. The marriage decline raises many social concerns, as there are plenty of social problems as a result of marriage decline, especially among the poor, such as out-of-wedlock childbearing, lower health and happiness outcomes, and reduced social supports. This paper provides new perspectives in understanding and solving these problems. For instance, reducing the wage gap between high and low educated men, probably by improving the skills for low educated men or improving the education especially for men can be helpful in solving the marital decline among low educated women.

CHAPTER 3. MALE INCOME INEQUALITY AND FEMALE MARITAL DECISIONS IN CHINA

3.1 Introduction

Since the launch of the economic reform in 1978, China has experienced an impressive improvement in the standard of living. Nevertheless, it has also witnessed a large increase in income inequality. The magnitude of this increase, exceeding all comparable records in other countries, is highly concerning (Yang 1998). The influences of rising income inequality on public health and education, crime and social stability, and long-term macroeconomic development are widely discussed and emphasized in previous research (Thorbecke and Charumilind 2002). Researchers obtained similar results from China as they did from other countries. For example, rising income inequality can lead to a lower happiness level, more tobacco, alcohol and drug consumption, higher mortality, lower sector growth as well as less macroeconomic stability in China (Zhao 2006; Li and Zhu 2006; Smyth and Qian 2008; Qin et al. 2009). However, the influence of income inequality on family formation in China seems to have been neglected.

In this paper, I focus on the relationship between income inequality and family formation in China. More specifically, I study the effect of male income inequality on young women's propensity to marry. I use this study to provide new evidence on the influences of rising income inequality in China.

The determinants of marriage formation have been widely discussed in the literature, such as women's economic independence, difficulty in transitioning to marriage, and the cost and benefit of the marital search process. Male earning inequality affects marriage formation in marital search theory. Rising wage inequality reflects the dispersion in the wage distribution of potential husbands, on which females are searching for males. The empirical evidence of the relationship between economic inequality and family formation has just been identified in recent years. Wage inequality is found to be a potential determinant of marriage trends in U.S (Stevenson and Wolfers 2007; Loughran 2002; Gould and Paserman 2003; Coughlin and Drewianka 2011). Rising male wage inequality is responsible for almost 1/5 of the decline in the marriage rate among 20 to 30 years old women from 1970 to 1990 in the U.S. (Loughran 2002; Gould and Paserman 2003).

Although the negative effect of male wage inequality on the marriage rate is convincing using U.S. data, no similar evidence has been found in other countries. As the socioeconomic backgrounds in other countries change, the results change. For instance, marriage is of more importance in Asian countries traditionally, as well as in many developing countries without an adequate social security system. Kuo (2008) investigates this link using data from Taiwan, but finds no significant effect of male wage inequality on female marriage propensity. This paper explores the evidence from China, an example of an Asian developing country, to add to the related studies.

Unlike previous literature, my research finds that rising male income inequality increases 20 to 34 years old Chinese women's propensity to marry. In each given year, the probability of women ever being married is higher in the city or county containing higher male income inequality. And within each city or county, the probability of women ever being married is higher in the year with higher male income inequality. The results are not caused by any observable individual characteristic of the woman, such as her age, education or income. They are not driven by the observed city or county characteristics such as sex ratio, average income level or the income gap between men and women. My results are obtained from five specifications, which I control for different levels of fixed effects and time trends. For example, in the most restrictive specification, I control for the year fixed effect and city/county fixed effect, as well as the national year trend. Almost all estimation results indicate a significant positive effect of male income inequality on female marriage propensity.

The results are robust with different variable measurements and subsamples. They appear to be stronger in certain subsamples than in the others. For example, the positive effects of rising inequality on marriage are significant for both urban and rural women, but it becomes more pronounced for rural women. It is no longer significant for subsamples of urban women under age 28 or urban women with at least a high school diploma.

The magnitude of these effects is not trivial. In the urban cities, a 10% increase of income inequality from the mean increases the 20 to 34 women's propensity to marry by at least 1.4 percentage points. The average propensity in samples is about 60 percentage points. The effect in rural counties is even bigger. A 10% increase of income inequality from the mean leads to a 3.5% higher propensity to marry for rural women aged 20 to 34, the sample average of which is about 63 percentage points. According to Yang (1998), the annual change of the overall Gini coefficient in China has been about 2.3% over 1981 to 1995. It can be assumed that a similar change happens in a typical city or county. For every 1000 women under age 34, 3 more urban women and 8 more rural women will get married in response to the higher income inequality each year. Taking into account the large population and the continual increase of income inequality in China, the effect cannot be neglected.

I provide two possible explanations to the results. The first one comes from a "marital search theory". Consider one dimension of marital search, when a woman searches for a husband according to the wage of men, conditional on other characteristics of the potential husband. On one hand, as the distribution of male wage becomes more dispersive and possibly fatter tailed especially on the upper side, women delay marriage because they form a higher standard on the minimum level of acceptable wage from their potential husband, which is called the "reservation wage". On the other hand, it is possible that a fatter tail on the upper side promotes marriage because it increases the possibility for a woman to meet someone above a given value of "reservation wage". Theoretically, when the male wage distribution has little mass in the center compared to that in the tails, the promotion effect is going to dominate. Otherwise, the delay effect is going to dominate. Empirical results suggest the delay effect is the dominant effect in the U.S. However, the promotion effect can be the dominant effect in China with polarization in the income distribution.

Another possible explanation comes from the belief that higher income inequality can lead to higher income risk and financial uncertainty. Since the family has risk sharing and social security functions, it becomes more desirable when income inequality increases. In line with this mechanism, I find that not only male income inequality increases female marriage propensity, the income inequality among all people also affects both men's and women's marriage propensity.

The results obtained in this paper have important implications. In general, they show that income inequality can raise the demand for families. However, the realities of marriage are not optimistic and are driven by many other factors. For instance, many men stay single especially those in poor rural areas, because of the high sex ratio of men over women and the out-marriage of single women, meaning rural women migrate to other villages to get married (Dus Gupta, Ebenstein, and Sharygin 2011). Meanwhile, the divorce rate in China has been rising continually.[17] In other words, marriage is more desirable under high income inequality and inadequate social security, but it is harder to form and easier to dissolve due to other factors. Hence, policies to reduce income inequality, to improve social security systems and to promote and strengthen marriage are necessary in China.

The rest of this chapter is arranged as follows: In section 3.2, I discuss the theories about marriage formation and the theories to link income inequality and marital decision. In section 3, I provide the Chinese background and studies of marriage formation. In section 3.4, I introduce the CHNS data and the estimation model. In section 3.5, I present the estimation results and the robustness checks as well as the discussion of different mechanisms. In section 3.6, I provide a conclusion based on the previous sections of this Chapter.

3.2 Literature Review

3.2.1 Determinants of Marital Decisions

The determinants of whether and when women decide to marry have long been studied. Becker (1973) first analyzed marriage from an economic perspective. He viewed marriage as a utility maximization choice determined by personal preferences and marriage market conditions. Marital gains come from the household specification between husbands and wives (Becker 1981).

Accordingly, many researchers look at the effect of women's economic independence featured by higher education, more labor force participation and higher wage on their propensity and timing to marry. Some argue that women's economic independence reduces their marital gains and increases the cost of household production and childbearing, hence causing marriage delay or decline (Oppenheimer 1988; Okun 2001). Others argue that women's higher earnings may promote marriage since marital gains also include risk sharing and consumption complementarities (Hess 2004; Stevenson and Wolfers 2007). It is still controversial whether women's economic independence can cause marital delay or decline (Santow and Bracher 1994; Oppenheimer 1994, 1997; Goldstein and Kenney 2001). However, researchers have reached a consensus that more labor force participation and higher earnings of men can raise the net benefit from marriage for both husbands and wives, hence promoting marriage (Pamela and Manning 1997; Xie et al. 2003).

Along with the discussion about marital gains, there is also some work focusing on the difficulties in the transition to marriage. There can be difficulties ranging from financial obstacles to school enrollment or military service (Oppenheimer 2003). Longer school enrollments are found to delay the marriage of young women (Goldstein and Kenney 2001). Due to these transition difficulties, cohabitation becomes a popular choice, which is facilitated by the sexual revolution and birth control development (Akerlof, Yellen, and Katz 1996; Goldin and Katz 2002).

Recent studies usually view marriage formation as a marital search process (Keeley 1977). From the women's perspective, they search over men with different personal characteristics. They make marital decisions based on their expected net benefit during each search period. The

[17] National Bureau of Statistics of the People's Republic of China, *Table 23.24 Number of marriages and divorces*, http://www.allcountries.org/china_statistics/23_44_number_of_marriages_and_divorces.html (2005)

net benefit is jointly determined by personal life-cycle preferences and marital market structures. According to Becker (1974), when the value of any expected improvement in the mate is no greater than the time and other inputs into the additional search, the search process will end. Let us call the "minimum level of utility that is acceptable to the woman or her family in a match" the "reservation level" (Montgomery, Cheung , and Donna 1988). Marriage occurs when the offer is available and the value to marry exceeds the "reservation level". Under this framework, factors that increase the offer rates and lower the reservation level can increase the possiblity of marriage at a given time. For example, higher sex ratio between men and women provides higher offer rates in the marriage market to women, hence increases the marriage rate of women (Schoen 1983; Angrist 2002).

3.2.2 Male Wage Inequality and Female Marital Search

Marital search theory provides a mechanism to explain the effect of male income inequality on female propensity to marry in China. Loughran (2002) links male wage inequality to female age at first marriage under the female marital search model, based on the job search model in Mortensen (1986). He assumes women search over male wages, denoted as w, conditional on the other observable characteristics of men, such as age, race, religion, appearance and education, denoted as H. At each period t, a woman gets at most one randomly drawn wage offer from a potential husband with probability q from the conditional wage distribution, denoted as $F(w|H)$. Let q be solely determined by the exogenous factors of the marriage market, such as the sex ratio. Let $W(w_t)$ be the present value of accepting the offer and U_t^s be her utility of being single in period t. Let c_t be her cost to pursue the offer in period t regardless of whether she gets the offer.

In this search problem, there exists a wage offer w^*, named the "reservation wage", such that the woman is indifferent from receiving the current offer and proceeding with an additional search in the next period. Let β be the discount factor. As shown in equation (3.1), at the reservation wage level, the net benefit of being married equals the discounted net benefit of searching in the future. In this period, the woman stops searching and accepts the current offer, and then gets married. Loughran (2002) further specifies the probability of getting married in equation (3.2), which is the product of the probability of getting an offer q and the probability of accepting the given offer $[1 - F(w^*)]$.

(3.1) $W(w^*) - U^s = \beta q \int_{w^*}^{\overline{w}} [W(w) - W(w^*)] dF(w|H) - c$

(3.2) $p = q[1 - F(w^*)]$

Two competing effects can occur as the male wage inequality enlarges in the search model. Assume the wage distribution can be fully depicted by its mean μ and standard deviation σ. Basically, keeping the mean constant, the increases of wage inequality, a higher σ, reflects the spread out of male wage distribution. It increases the reservation wage w^*, thus lowering the possibility to accept a given offer $[1 - F(w^*)]$. We shall call this the "marital delay effect". Meanwhile, the spread of wage distribution also increases the possibility to meet someone above the reservation wage $[1 - F(w^*)]$, which is the possibility to accept the offer at any given value of w^*. We shall call this the "marital promotion effect".

The key to determining which effect on $[1 - F(w^*)]$ dominates is the shape of the distribution. As mentioned in Gould and Paserman (2002), when the distribution has more mass in the center than the tails, the delay effect dominates. For example, in Figure 3.1, as σ enlarges, wage distribution changes from $F_1(w|H)$ to $F_2(w|H)$. The reservation wage increases from w_1^* to w_2^*. The area above w_1^* in distribution $F_1(w|H)$ is bigger than the area above w_2^* in distribution $F_2(w|H)$. On the contrary, when the distribution has more mass in the tails than in the center, the promotion effect dominates (Burdett and Ondrich 1985). For example, in Figure 3.1, as σ enlarges, the wage distribution changes from $F_2(w|H)$ to $F_3(w|H)$. The reservation wage increases from w_2^* to w_3^*, $F_3(w|H)$ has more mass in the tails than in the center and the area above w_2^* in distribution $F_2(w|H)$ is smaller than the area above w_3^* in distribution $F_3(w|H)$.

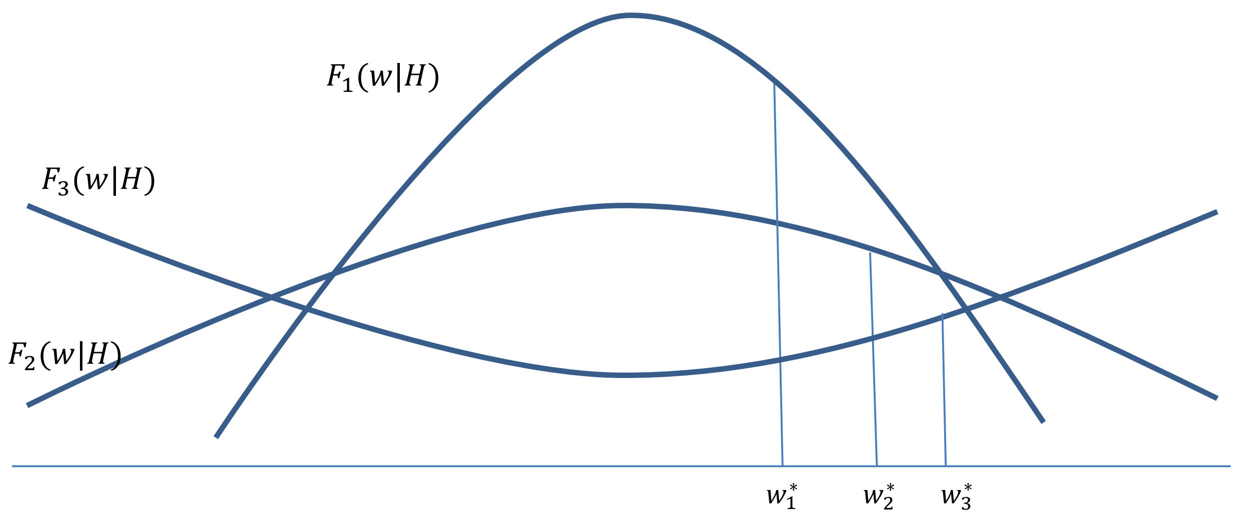


Figure 3.1 The Probability of Accepting A Marriage Offer Under Different Male Wage Distributions

3.2.3 Marry to Be More Secure

Another mechanism could be that the income risk and social uncertainty caused by higher income inequality will raise the demand for marriage with social security functions. People get mutual supports, elderly care, and income risk-sharing over the lifetime from family members. Becker (1974) argues that the creation of one's own children is the fundamental reason for marriage. Elder care heavily relies on spouses and children in developing countries like China. The desire to share income risk promotes marriage (Hess 2004). When income shocks are stronger and more frequent and social uncertainty is higher, it becomes more appealing to form families, especially in countries that lack adequate social security systems.

The changes in male income inequality may just reflect changes in overall income inequality. Overall income inequality raises the financial risk and uncertainty of the society (Gottschalk and Moffitt 1994). For example, changing income distribution is associated with higher income mobility. Income inequality is also associated with social instability, such as higher crime rates (Choe 2008). In general, higher income inequality leads to higher risk and uncertainty economically and socially, which leads to a higher demand for marriage. The effects may vary with different causes of income risk. Firstly, it is possible that income mobility is mainly within groups defined by education, experience or age (Geweke and Keane 2000; Schmidt 2008). In this case, it is generally more beneficial for poor people, both men and women, to get married with richer partners. Specifically, people with different risk tolerance could also behave differently. For instance, risk neutral or adverse women/men usually want to marry up with rich men/women in order to obtain a better financial condition (Hess 2004; Coughlin and Drewianka 2011). However, high risk tolerant women/men may tend to delay marriage in pursuit of their own financial stability and to make a proper choice of their future husband/wife (Oppenheimer, Kalmij, and Wackero 1997; Schmidt 2008). Secondly, it is possible that higher income mobility comes from between-group inequality, because that the same individual's income may rank differently in the income distribution over time (Gottschalk 1997; Gottschalk and Moffitt 1994; Moffitt and Gottschalk 2011). As a result, both young men and women want to share income risk over lifetime with their spouses. With the exception that younger women may wait longer for a more credible signal from men and marry later, since younger men tend to have higher earning mobility than older men (Buchinsky and Hunt 1999).

In conclusion, competing effects exist when income inequality affects marriage formation. The effects vary for people with different demographic features. It is not clear to determine, however, the overall impact of income inequality on marriage formation only by checking the theories.

3.3 Background

3.3.1 Policies

Individuals' marital timing has been affected by policies in modern China. In the early 70s, population policy encouraged young people to delay marriage (Ye 1992). In 1980, the New Marriage Law sets the "minimum age to marry" to be 22 years for men and 20 years for women (Xia and Zhou 2003). Han (2010) explains the median age at first marriage between 1970 and 2000 in China by policies. Accordingly, the initial drop of age at first marriage in the 70s and early 80s was a reflection of the "late marriage" policy; the steady increase from the mid-1980s to 2000 was due to the lower legal marriage age and the New Marriage Law in 1980. The delay trend of marriage from 1985 to 2000 was also the result of education reform.

The choices of partners have also been affected by polices. Since 1950, arranged marriage has been deemed illegal by the Marriage Law. From 1966 to 1978, the "Cultural Revolution" encouraged young adults to marry people outside their own educational groups especially to marry those in the working class, such as workers, farmers and soldiers. Since the end of the "Cultural Revolution" in 1978, the choice to select partners became freer (Xia and Zhou, 2003). Accordingly, the education homogamy rates, i.e. the rates of people marrying with similar education levels, experienced a decrease from 1970 to 1980, a substantial increase from 1980 to 1995 and a slower growth in late 1990s (Han 2010).

Concurrently, the marriage market has been influenced by policies. Since 1978, the "reform and opening-up" policy has resulted in a major improvement in economical and educational outcomes. Women's social standing has been stepped up and their educational attainment has caught up with men's (Mu and Xie 2011). The "one-child policy"[18], starting in 1978, has resulted in an unbalanced sex ratio in China, mainly due to the preference of sons, especially in rural areas (Zhu, Lu, and Hesketh 2009). During the reforms for state-owned enterprises in the mid-1990s, a massive number of workers were laid off, and urban labor force participation was

[18] According to "one child policy", urban couples can have one child without fines. Rural couples, couples with at least one ethnic minority and couples both with no siblings can have a second child without fines.

reduced by almost 10%. Also noteworthy is that more females were laid off at all working ages than males (Maurer-Fazio, Hughes, and Zhang 2005).

3.3.2 Income Inequality, Polarization, and Mobility

Along with economic reform, income inequality has also been increased. From 1978 to 2006, the overall Gini coefficient had increased from 0.31 to 0.45; the rural Gini coefficient changed from 0.21 to 0.37 and the urban Gini coefficient changed from 0.16 to 0.34 (Yao and Wu 2010). The wage gap between males and females has increased; the urban/rural economic gap has widened; and the regional inequality between eastern provinces and the interior and western provinces has grown (Maurer-Fazio et al. 2005). Income inequality has increased both within and between rural counties between 1988 to 1995 (Gustafsson and Li 2002).

Polarization has been rising together with income inequality exaggeration in China (Chen and Zhang 2009). Kanbur and Zhang (2001) varified that the overall trends of income inequality and polarization were rising from 1983 to 1995, by using three newly developed measures of polarization and two traditional measures of income inequality. Meanwhile, income mobility has become higher for the poorest 25% of the people, but lower for the middle class people, using rural data from 1987 to 2002 (Zhang, Huang and Mi 2006).

3.3.3 Marital Propensity and Timing

Studies on marriage formation in China are limited; among which no papers have linked income inequality to marriage formation. Existing papers find some similar features as in western countries. For instance, urban people usually marry later than rural people and men usually marry later than women (Li 1985; Ru, Lu, and Li 2007). Better education and higher regional GDP growth rate reduces the tendency for young adults to get married. Urban (but not rural) men and women with higher wages tend to delay marriage (Xu, Qiang, and Wang 2003).

Researchers have paid attention to the problems due to the unbalanced sex ratio in China. The significant unbalanced sex ratio is mainly caused by sex selective abortion under the “one child policy”. It is worse in poor rural areas due to women’s “out-marriage”, when they migrate out of the village to get married (Meng 2009). As predicted by Dus Gupta, et. al (2011), “one in five men will fail to marry in 2020” without the proper policy and this will mostly happen to men in the rural areas of poor provinces.

3.4 Data and Model

3.4.1 Data

The micro data used in this paper is from the China Health and Nutrition Surveys (CHNS). It is conducted by the Carolina Population Center at the University of North Carolina and the National Institute of Nutrition and Food Safety at the Chinese Center for Disease Control and Prevention. The survey started in 1989 and repeated in 1991, 1993, 1997, 2000, 2004, 2006 and 2009. Each year it drew samples from about 4400 households that contained around 19,000 individuals. They were located in nine provinces, namely the Jiangsu, Shandong, Heilongjiang, Liaoning, Henan, Hubei, Hunan, Guangxi and the Guizho.[19] These provinces varied substantially in geography, economic development, public resources, and general health indicators. Table 3.1 denotes the region and GDP per capita in 2007 of each province.

A multistage, random cluster process was used in each province. Two urban cities in each province were surveyed, which included the provincial capital and another large city. Four rural counties[20] in each province were randomly selected by a weighted sampling scheme. Urban and suburban neighborhoods within the cities and villages and townships within the counties were randomly selected.[21]

Three datasets were merged, which respectively contained demographic, education, and income features, into one unbalanced panel. The merged dataset has a total of 103,764 observations, among which 51,775 are men. Each observation is uniquely identified by a household ID, a household member ID, and the survey year. In total, 27,783 individuals had participated in the survey at least once. Only 2,952 among them had participated in all eight waves of surveys.[22]

Table 3.1 Provinces in CHNS Survey, Region and GDP Per Capita

Provinces	Region	2007 GDP Per Capita
Jiangsu	East	32,985
Shandong	East	27,148
Liaoning	Northeast	24,645
Heilongjiang	Northeast	18,463
Henan	South Central	15,056
Hubei	South Central	14,733
Hunan	South Central	13,123
Guangxi	South Central	11,417
Guizhou	Southwest	6,742

Notes: GDP per capita (Yuan). Reprinted from World statistics yearbook in 2007 (p.30-32), by Wang et al, 2007, Beijing: China Financial and Economic Publishing House

3.4.2 Framework

According to previous literature, the decision to marry in a certain year for a woman is related to factors that indicate her economic independence and the difficulties in transitioning to

[19] Heilongjiang was not surveyed in 1989, 1991 and 1993. Liaoning was not surveyed in 1997. Other provinces were surveyed all eight years.

[20] Counties are geographic units within rural areas in each province.

[21] The survey does not provide representative weight value for each individual. To control for multistage sampling and an array of multilevel modeling issues, it is recommended to adjust standard errors by clustering at the community level if possible, according to the project senior programmer.

[22] 5,765 individuals had participated only once, 4,183 twice, 4,736 three times, 3,497 four times, 3,544 five times, 1,381 six times, and 1,725 seven times.

marriage. These can be captured by the woman's personal attributes such as age, wage and education levels. Again, according to search theory, the marriage market features also affect a woman's marriage decision-making. Sex ratio affects the offer rates q as in equation (3.2). The mean income level and gender wage gap affects the income distribution F as in equation (3.2). And the key factor I am interested in is the income inequality of males, which affects the possibility to accept a marriage offer, the value of $[1 - F(w^*)]$ as in equation (3.2). The marriage decision is also shaped by unobserved changes in the marriage market, such as related policies and socioeconomic trends.

To summarize, the decision to marry for woman (i) in city/county (j) in year (t) is a function of male income inequality, $Ineq_{jt}$, her personal attributes, $Attr_{ijt}$, and other observed, M_{jt}, and unobserved, U_{jt}, marriage market conditions, as noted in equation (3.3). Altogether, the reduced form model is stated in equation (3.4).

(3.3) $P_{ijt} = F(Ineq_{jt}, Attr_{ijt}, M_{jt}, U_{jt})$

3.4.3 Baseline Specification

(3.4) $P_{ijt} = a * Ineq_{jt} + b' * Attr_{ijt} + \mathrm{c}' * M_{jt} + U_{jt} + \varepsilon_{ijt}$

In the baseline estimation, I test a linear probit model as denoted in equation (3.4) using the OLS method. P_{ijt} denotes the marital status of an individual female (i) in city/county (j), which contains the value of one if the woman has ever been married in year (t) and zero otherwise; $Ineq_{jt}$ is the measure of male earning inequality within a city/county in a year; $Attr_{ijt}$ is a vector of the female personal attributes, such as age, education and income; M_{jt} is the vector of observable marriage market features, such as the sex ratio, average male income, and gender income gap; U_{jt} is a serial of controls for the unobserved year and geographic fixed effect as well as the time trends. The error term ε_{ijt} follows a standard normal distribution, N(0,1) . The key coefficient is (a). If (a) is positive, woman (i) is more likely to marry in a city/county (j) with a higher male income inequality in year (t).

I employ marital status in the surveyed year as the dependent variable.[23] Included women are aged from 20 to 34, which is the prime age range for females to marry and to reproduce. To be specific, dummy variable P_{ijt} in equation (3.4) equals one if the woman has ever married, and zero if she has never married in year t.[24] Since a married woman remains married in the sample even with a change in the male wage inequality in the city or county over time, keeping the repeated sample of married women in the estimation will cause a bias. Therefore, I delete repeated married samples. As a result, a woman will only appear once as "ever married" in the estimation sample. For example, a woman married in 2004 will respond as "never married"

[23] "Age at first marriage" is a clear measure for marital timing, but the question about "age at first marriage" was only asked in the first wave of the surveys. Even given the "age at first marriage", it is impossible to match the information of the marriage market when marriage actually occurred with the woman in the sample within the same dataset. Besides, I could not obtain the names of the rural counties and the other big city besides the capital from the public data. Hence I could not match it with other datasets.

[24] There are five possible choices, respectively "1" never married, "2" married", "3" divorced, "4" widowed, and "5" separated. Women whose answer is among 2 to 5 are classified as "ever married".

before 2004, and “ever married” in 2004. Any observations of this woman after 2004 will be deleted.

The main explanatory variable is the income inequality among 20 to 50 years old males. Total income is probably a more important factor than wage in marital search. If using wage inequality only, as in other literature, sample size will become much smaller especially for rural estimation. In the baseline regression I use the Gini coefficient calculated by the income of 20 to 50 years old males in a city or county in each year.[25] Income is calculated by the total real individual yearly income in the previous year.[26] It includes all income sources from business, farming, fishing, gardening, livestock, non-retirement wages and the retirement wages.

I control for personal characteristics including age and education level in the baseline estimation. I categorize the age variable into seven small blocks, and create one age dummy for each block, due to the small sample size at each age. For example, there is one age dummy for women who are aged between 20 to 22 years except for exactly 22, one for those aged between 22 to 24 years except for exactly 24, and so forth. I create an indicator for education level according to the highest degree obtained by the respondents up to the survey year.[27] The indicator value is 3 if the woman has postsecondary degree, including technical or vocational, university or college, and master or higher degree; 2 if she has high school diploma; and 1 if she has a middle school education; and 0 if she has no education at all. I define a marriage market as the group of prime age men and women in a city or county in one year.

There are other personal attributes such as wage and ethnicity to be considered. The female wage or labor force participation could be added in order to control for the “female independence effect” on marriage propensity based on the literature review. However, the missing values of such variables can lead to further attrition of samples. By excluding females without work, the test also loses the comparison between working women and non-working women. Therefore, I choose to exclude these variables in the baseline regression so that I could keep both working and non-working women in the sample. According to the later robustness check, further controlling for the female wage or income does not change the basic results of the baseline estimation. Moreover, I do not control for ethnicity, as the number of women reported to be in the national minority is only 112. In addition, I cannot detect any statistical difference in marital status in the sample between women of a national minority and women of the Han ethnic group.

I control for the observable marriage market conditions including the average male income level, the income gap between genders and the sex ratio. They are constructed by the original samples before dropping the repeated “ever married” observations. Average male income is the arithmetic means of the log income values of all men aged from 20 to 50 in urban areas. Mean gender income difference is calculated by the average male income minus average female

[25] I only use males’ income within this dataset to create inequality of income between males, because I cannot identify every city or county in order to match with other datasets.

[26] The survey asks questions about individual’s income or wage in last year.

[27] The options of question for education levels are “missing, 0 for none, 1 for graduated from primary school, 2 for lower middle school degree, 3 for upper middle school degree, 4 for technical or vocational degree, 5 for university or college degree, 6 for master or higher”. Here, primary school usually takes 6 and sometimes 5 years to finish; lower middle school takes 3 years; upper middles school also named high school takes 3 years; technical or vocational degree is professional training for students who don’t go to college. The compulsory education requires young people to finish primary school and lower middle school education.

income aged from 20 to 50. The sex ratio is calculated by the number of males divided by the number of females in the age range of 20 to 34 in each marriage market.

There are still unobserved marriage market features that may correlate with both the individual's marital choice and the earning inequality. I control for these unobserved factors by adding city or county fixed effects, national trends, provincial trend, and time trend within a city in urban area or a county in rural area. I add dummies for the cities (or counties) to control for city (or county) fixed effect. I further add dummies to indicate the province capital city for urban estimation. I use a dummy variable for each wave of the survey as a year fixed effect. Moreover, I use the value of the difference between the surveyed year and 1989 as year trends, the product of year trend and province dummies as province specific time trends, and further the product of year trend and city/county dummies as city/county specific time trends.

Finally, variables defined in a marriage market are in a higher aggregation level, say the city-year level, than individual units. Meanwhile, there is a grouped structure of errors within a city in a year (Gould and Paserman 2003). Hence, I cluster standard errors by city-year unit. There are 136 clusters in every specification included in the baseline estimation.

In the baseline estimation, there are 6,834 female observations with 2,241 urban female observations and 4,593 rural female observations. They are collected from 1,807 urban women

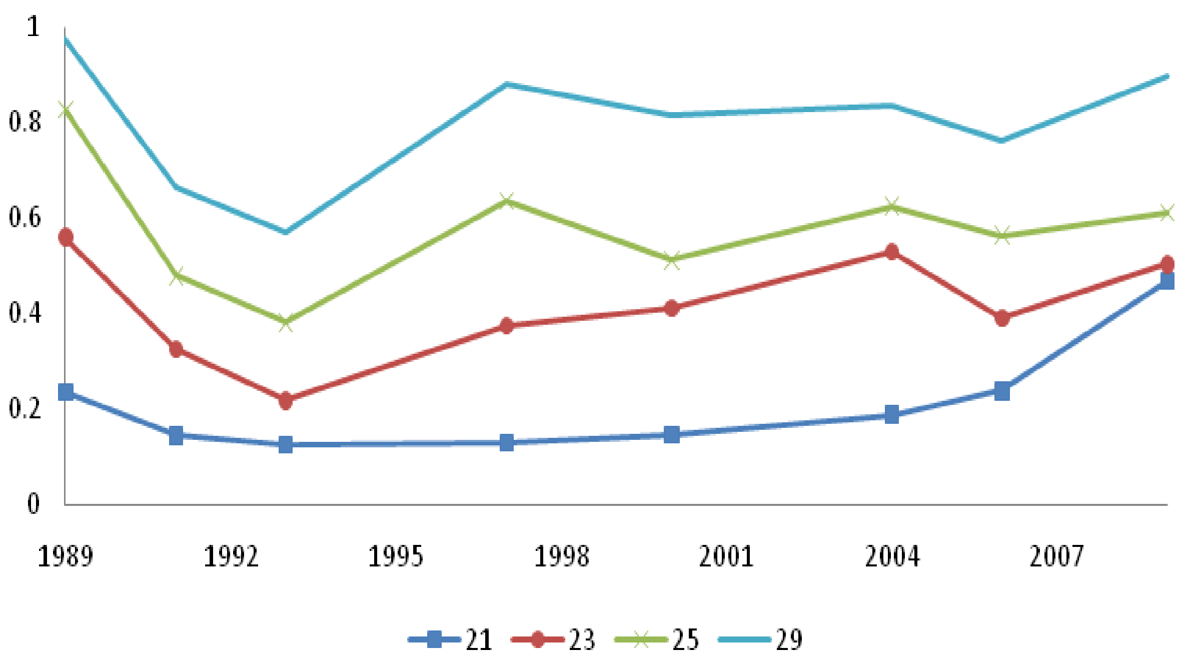


Figure 3.2 Proportions of Married Women by Age and Year in Samples for Baseline Estimation

Notes: Each dot represents the proportion of married women among all women in a certain age range in a year. Specifically, the value is calculated by dividing the number of married women (who are married before current survey but after last survey or newly added married women) by the number of women (never married before last survey, or newly added) within the age range of 20 to 22, 22 to 24, 24 to 26 or 28 to 30 in each surveyed year. Each age specification has eight values from the eight waves of the surveys, respectively 1989, 1991, 1993, 1997, 2000, 2004, 2006, and 2009 in the baseline sample. I exclude the dot where the number of women in the defined age group and year is smaller than 50. This leads to 2, 4 and 4 missing dots respectively for age around 27, 31 and 33. Hence, the lines of those age range are not shown.

and 3987 rural women. Figure 3.2 shows the proportion of married women within each age-year combination. Within each year, the marriage rate generally increases as age increases. The marriage rate around each age tends to decrease between 1989 and 2003, but tends to increase after 2003.

Table 3.2, part A, lists means of selected variables by year in the urban sample. Sample

Table 3.2 Sample Means of Women Aged 20 to 34

Variables	Year							
	1989	1991	1993	1997	2000	2004	2006	2009
A. Urban								
Percentage of married women	0.77	0.31	0.21	0.63	0.48	0.58	0.43	0.57
Age	26.61	23.33	23.38	26.42	25.52	26.65	29.23	26.29
Percentage of illiteracy	0.10	0.06	0.04	0.06	0.02	0.02	0.02	0.02
Percentage with only compulsory	0.53	0.54	0.52	0.51	0.38	0.34	0.26	0.27
Percentage of high school graduates	0.25	0.26	0.29	0.23	0.22	0.23	0.16	0.18
Percentage with postsecondary degree	0.11	0.14	0.15	0.20	0.39	0.40	0.56	0.54
Std of male log income	0.75	0.71	0.76	0.84	0.86	0.91	0.81	0.87
Mean male log income	8.07	8.21	8.42	8.69	8.92	9.17	9.55	9.74
Mean female log income	7.91	8.02	8.29	8.42	8.71	8.92	9.11	9.48
Gini of male income	0.41	0.33	0.39	0.40	0.37	0.37	0.39	0.38
Sex ratio (20-34)	1.01	1.00	1.00	1.00	0.98	0.95	0.94	0.96
Sex ratio (20-50)	0.98	0.96	0.97	0.98	0.97	0.92	0.89	0.91
Number of observations	658	205	178	380	269	215	148	188
B. Rural								
Percentage of married women	0.76	0.30	0.26	0.56	0.50	0.69	0.65	0.73
Age	26.43	22.79	23.05	25.31	24.79	25.95	29.28	25.65
Percentage of illiteracy	0.20	0.16	0.10	0.08	0.02	0.05	0.04	0.02
Percentage with only compulsory	0.61	0.68	0.67	0.72	0.69	0.70	0.66	0.63
Percentage of high school graduates	0.14	0.11	0.14	0.11	0.14	0.12	0.12	0.14
Percentage with postsecondary degree	0.05	0.05	0.09	0.10	0.15	0.14	0.18	0.21
Std of male log income	0.99	0.86	0.93	0.93	1.00	1.06	1.05	1.08
Mean male log income	7.87	7.87	7.95	8.31	8.53	8.62	8.98	9.43
Mean female log income	7.58	7.70	7.74	8.09	8.20	8.27	8.53	9.08
Gini of male income	0.46	0.41	0.42	0.41	0.44	0.44	0.41	0.48
Sex ratio (20-34)	1.01	0.99	1.00	0.99	1.00	0.96	0.96	0.96
Sex ratio (20-50)	1.01	0.99	1.01	1.03	1.05	0.94	0.91	0.93
Number of observations	1,311	493	466	792	659	326	242	304

Notes: The income is a constructed individual yearly income in the previous year (Yuan) inflated or deflated to the 2006 value. The Gini coefficient of male income, standard deviation and mean of males' log income are calculated from the income of males aged from 20 to 50 within each city in the survey year. Means of female log income are calculated from the income of females aged from 20 to 50 within each city in the survey year. Values (except for the number of observations) shown are the mean value of variables averaged by females aged from 20 to 34 in each wave of the surveys.

means of education levels reflect a significant improvement in education of young urban females over the last 20 years. The average male and female log income levels are increasing over years. I calculate two sex ratios for men and women who are within the age range of 20 to 34 and 20 to 50 respectively. The mean values of sex ratio tend to be smaller than typical results in other studies (Zhu et al. 2009). It may be because that it is not a national representative sample and the sample size in each city during each year is limited. Table 3.2, part B, shows the variable means for the rural sample. With a larger sample size, the means of sex ratio are higher than urban results and become closer to typical results. Nevertheless, the degree of the variation over time and across cities can still be informative.

3.5 Results

I estimate the urban and rural samples separately using the baseline model (3.4). I adopt five specifications: (1) controls for year fixed effect, national time trend, province and capital fixed effect; (2) controls for city fixed effect and national time trend; (3) controls for city fixed effect, national time trend and province-specific time trends; (4) controls for city fixed effect, national time trend and city-specific time trends; and (5) controls for year fixed effect, city fixed effect and the national time trend.

3.5.1 Urban Baseline Estimation

For urban samples, in the first specification, I test the effect of cross-city variation of male income inequality on young female marriage propensity. In addition to individual attributes and marriage market features, I control for national time trend, year fixed effect, province fixed effect, and the province capital fixed effect. National time trend controls the economic, social or cultural trends in China over time, such as the long run economic growth and improvement of education level. Year fixed effects account for the unobserved changes in each year which affect all cities, such as national policies that affect marriage formation. Province fixed effect control for unobserved differences among provinces, such as special custom for marriage and different economic environments, and policies varying across provinces. The dummy variable to indicate whether the city is the capital city accounts for the unobserved differences between capital and the other large city in the same province.

The results are shown in column (1) Table 3.3. Marginal effects of selected variables are listed with robust standard errors. The Gini coefficient has a significant positive impact on the propensity to get married for women aged 20 to 34 from 1989 to 2009. The marginal effect has a value of 0.366 at the 5% significant level. This means that when the Gini coefficient used to measure male income inequality in a city becomes 0.01 higher than the mean value, which is around 0.386, across all sampled cities, the probability to be married for women aged 20 to 34 will be 0.366 percentage point higher than the mean value, which is around 60 percentage points. The absolute value of the marginal effect decreases with age. The younger the women are, the less likely for them to be married. The propensity to marry before 30 is negative and significant at the 1% level, compared with the women around 33 years old. Women who have finished about 9 years of compulsory schooling (middle school graduate degree) have a higher probability to be married than the illiterate young women by a marginal effect of 0.148. That is to say, women who have completed the compulsory schooling are 14.8% more likely to get married

Table 3.3 Probability of Being "Ever Married" for Urban Women Aged 20-34 (1989-2009)

		(1)	(2)	(3)	(4)	(5)
Dependent variable: 0, never married; 1, ever married.						
Gini of male income		0.366** (0.181)	0.604*** (0.233)	0.637*** (0.244)	0.602** (0.267)	0.352* (0.187)
Education	Illiteracy	–	0.161*** (0.056)	0.184*** (0.055)	0.184*** (0.056)	0.155*** (0.056)
	Compulsory education	0.148** (0.064)	0.297*** (0.036)	0.310*** (0.036)	0.317*** (0.035)	0.307*** (0.036)
	High school graduates	-0.017 (0.065)	0.135*** (0.032)	0.144*** (0.032)	0.148*** (0.032)	0.147*** (0.031)
	Postsecondary Degree	-0.169** (0.068)	–	–	–	–
Age	20≤age<22	-0.763*** (-0.023)	-0.771*** (0.023)	-0.774*** (0.023)	-0.740*** (0.024)	-0.762*** (0.024)
	22≤age<24	-0.633*** (0.040)	-0.657*** (0.037)	-0.660*** (0.038)	-0.594*** (0.037)	-0.630*** (0.041)
	24≤age<26	-0.476*** (0.059)	-0.500 (0.055)	-0.507*** (0.056)	-0.400*** (0.050)	-0.471*** (0.060)
	26≤age<28	-0.312*** (0.076)	-0.322*** (0.074)	-0.328*** (0.075)	-0.206*** (0.060)	-0.306*** (0.077)
	28≤age<30	-0.209*** (0.075)	-0.232*** (0.074)	-0.245 (0.075)	-0.114* (0.061)	-0.203*** (0.075)
	30≤age≤32	-0.114 (0.083)	-0.129 (0.080)	-0.136* (0.082)	0.128* (0.072)	-0.103 (0.083)
Sex ratio (20-34)		-0.172 (0.163)	-0.137 (0.299)	-0.228 (0.304)	-0.810** (0.379)	-0.057 (0.210)
Mean male log income		-0.007 (0.078)	0.060 (0.096)	0.050 (0.102)	0.107 (0.102)	-0.001 (0.102)
Mean gender difference in log income		-0.021 (0.080)	-0.082 (0.105)	-0.046 (0.111)	-0.139 (0.130)	-0.025 (0.081)
Year Dummies		Yes	No	No	No	Yes
Year Trends		National	National	National& Province	National& City	National
Geographic fixed effects		Province& Capital	City	City	City	City
Number of city-year clusters		136	136	136	136	136
Number of observations		2,241	2,241	2,241	2,241	2,241

Notes: ***p<0.01, **p<0.05, *p<0.1. The coefficients are the marginal effects of explanatory variables evaluated at means on the probability of being married estimated from a probit model. The sample includes 2,241 female observations aged 20 to 34 collected from eight waves of surveys (1989, 1991, 1993, 1997, 2000, 2004, 2006 and 2009), which were answered by 1,807 unique women living in 18 urban cities including the province capitals of nine provinces (Heilongjiang, Liaoning, Shandong, Jiangsu, Henan, Hubei, Hunan, Guangxi and Guizhou) in China. Standard errors in parentheses are adjusted by city-year clusters.

than those who have no education. But women with a postsecondary degree would delay marriage with a marginal effect of -0.169. As a result, they are 16.9% less likely to get married compared with those illiterate women.

I control for the city fixed effect in the other four specifications shown in Table 3.3 column (2) to (5). I explore the effect of the variation of male income inequality over time, in a city, on a woman's marital propensity in that city. In the second specification, there are controls for city fixed effect and national time trend. The marginal effect of male income inequality measured by the Gini coefficient is 0.604 and significant at the 1% level. This means when the women in a city experience a higher male income inequality over time, which is a deviation from the national trend, they are more likely to be married. If the Gini coefficient for male income increases by 0.01, the women's propensity to get married increase by 0.604 percentage point compared with the mean of 60 percentage points.

In the third specification, I control for province-specific time trends in addition to all the controls in the second specification. The marginal effect is 0.637, which has a greater magnitude than that in the second specification without province-specific time trends. Hence, higher male income inequality in a city over time, which is a deviation from the provincial trend and national trend, may have positive effects on female marriage propensity. If the male income Gini coefficient is higher than the mean, say 0.386, by 0.01, then the probability to get married for women aged 20 to 34 becomes 0.637 percentage point higher compared with the mean of 60 percentage points.

In the fourth specification, I further control for the city-specific time trends. It only identifies the variation of the male income Gini coefficient over time that is a deviation from the city-specific year trend in each city conditional on the national year trend. The marginal effect is 0.602 and significant at the 5% level. Hence, if the male income Gini coefficient is 0.01 higher than the average value, the probability for a woman in that city to get married is 0.602 percentage point higher than the average probability, about 60 percentage points.

In the last specification, I add both city fixed effect and year fixed effect, as well as the national year trend. This is the most restrictive specification; it identifies the variation in the male Gini coefficient for a year, which is a deviation from the average level in that city. In this case, the marginal effect is 0.352 at the 10% significant level, shown in the column (5) of Table 3.3. If the male income Gini coefficient has a value 0.01 higher than the average, the marriage propensity for women in the city to marry is 0.352 percentage point higher than the average (about 60 percentage points).

To better interpret the magnitude of the marginal effect, I display the prediction results of the marginal effects using the delta method in Table 3.4, part A. I show the results by using different ways to change the value of the male income Gini coeffficent. I include the prediction for the results of the first four specifications, which are all significant at the 5%. For example, according to the result from the first specification in the baseline regression, a city with 10% higher male income inequality than the mean (mean*1.1) would lead to 1.4 percentage points higher propensity to marry among young females compared with other cities in the same year, where the average propensity to marry for a 20 to 34 years old woman, in a city, in the sample is 60.2 percentage points. According to the results from the fourth specification, when the male income inequality in the same city becomes one standard deviation bigger than before, which is 0.485 compared with 0.386, marriage propensity among women aged 20 to 34 in that city becomes 5.8 percentage points higher than the mean over time, which is 59.9 percentage points.

Table 3.4 Predictions from Baseline Results

	Predicted delta probability of being "ever married"			
	(1)	(2)	(3)	(4)
A. Urban	Sample mean probability			
	0.602	0.595	0.597	0.599
Mean Gini = 0.386	Difference from the sample mean probability			
Mean Gini+1*std =0.485	0.036	0.059	0.062	0.058
Mean Gini+2*std =0.584	0.070	0.114	0.119	0.115
Mean Gini*0.9 =0.347	-0.026	-0.023	-0.025	-0.024
Mean Gini*1.1=0.425	0.014	0.024	0.025	0.023
B. Rural	Sample mean probability			
	0.601	0.624	0.626	0.629
Mean Gini = 0.437	Difference from the sample mean probability			
Mean Gini+1*std=0.525	0.023	0.055	0.050	0.069
Mean Gini+2*std=0.613	0.039	0.107	0.099	0.133
Mean Gini*0.9=0.393	-0.012	-0.029	-0.027	-0.038
Mean Gini*1.1=0.481	0.011	0.028	0.025	0.035

Notes: The label column indicates the mean value of Gini coefficient and the value after different treatments. Column (1) to (4) reflects the specification (1) to (4) in the baseline regressions. The top row indicates the mean probability to marry. The following rows show the change from the mean probability due to the treatment.

3.5.2 Urban Robustness Check

In Table 3.5, I show the robustness checks of the marginal effect of male income inequality on 20 to 34 years old females' propensity to marry. As suggested by the survey, to control for the design effect, it is better to control for the urbanization ratio and adjust standard errors by communities. The income values are already adjusted by the urbanization ratios. Hence, I further adjust standard errors by community-year units. The coefficients are still significant with the same signs as in the baseline results.

I also investigate the impact with some alternative measures of the male income inequality. In previous literature, the standard deviation of male log wage is the main measure of male earning inequality. When I use this measure for males aged 20 to 50, the results are positive, but insignificant. This effect may be caused by a smaller sample size to calculate the standard deviation. It is also possible that the Gini coefficient can reflect some unobserved factors, but the standard deviation cannot. This possibility is examined in the discussion section below.

In addition to the standard deviation, I also use three other measures of the male income inequality, the 90/10 ratio, the 90/50 ratio and the 50/10 ratio. The 90/10 ratio divides the male income ranked at the 90th percentile by that ranked at the 10th percentile in the male wage distribution. The 90/50 ratio divides the income of the 90the percentile by the median. It focuses on the income inequality in the upper tail of the distribution. The 50/10 ratio, on the contrary, captures the income inequality in the lower tail of the distriubtion. The coefficients are very small and insignificant for most specifications, except that in the specifications (3) and (4) there are negative results using the 50/10 ratio. This indicates the time varying male income inequality in the lower tail deviated from the province-specific or city-specific year trends decreases the marital propensity of young females. It is consistent with the fact that fewer marriageable men lead to fewer marriage.

Table 3.5 Robustness Check of Probability of Being "Ever Married" for Urban Women Aged 20-34 (1989-2009)

Alternative measures		Marginal effect of male income inequality				
		(1)	(2)	(3)	(4)	(5)
Baseline results of male income Gini		0.366** (0.181)	0.604*** (0.233)	0.637*** (0.244)	0.602** (0.267)	0.352* (0.187)
Standard errors clustered in community-year		0.366* (0.194)	0.604*** (0.213)	0.637*** (0.228)	0.602** (0.259)	0.352* (0.202)
Alternative measures of male income inequality	Std of male log income	0.036 (0.083)	0.080 (0.100)	0.065 (0.095)	0.101 (0.097)	0.038 (0.081)
	90/10 ratio	0.000317 (0.000224)	-.0001378 (0.000345)	-0.000353 (0.000326)	-0.000490 (0 .000335)	0.000246 (0.000280)
	90/50 ratio	-0.00821 (0 .0181)	0.00450 (0.0234)	0.00126 (0.0232)	0.000904 (0.0239)	-0.00936 (0.0179)
	50/10 ratio	.000566 (0 .000613)	-0.00125 (0.000817)	-0.00190* (.000787)	-0.00226* (0.000787)	0.000237 (0 .000778)
Sex ratio	20-50	0.346* (0.178)	0.598*** (0.231)	0.629*** (0.241)	0.622** (0.257)	0.350* (0.185)
Average income	Mean income all 20-50	0.326* (0.190)	0.396* (0.237)	0.468* (0.261)	0.381 (0.281)	0.303 (0.200)
	Mean gender difference in income	0.258 (0.219)	0.429 (0.272)	0.521* (0.287)	0.493* (0.298)	0.242 (0.243)
Restrict samples	Female age 20-30	0.422** (0.186)	0.661*** (0.250)	0.666*** (0.260)	0.665** (0.284)	0.394** (0.198)
	Female age 20-28	0.276 (0.220)	0.266 (0.247)	0.219 (0.252)	0.116 (0.278)	0.141 (0.211)
	Female with education	0.434** (0.191)	0.630*** (0.239)	0.629** (0.249)	0.675** (0.267)	0.392** (0.195)
	Female with at least high school degree	0.329 (0.270)	0.313 (0.287)	0.163 (0.303)	0.126 (0.353)	0.189 (0.247)

Notes: ***p<0.01, **p<0.05, *p<0.1. Marginal effects of male income inequality on young urban female probability of being ever married are shown. Column (1) to (5) is respectively consistent with the specification (1) to (5) in the baseline regression. Standard errors are corrected by city-year (or community-year for row 2 only) clusters shown in parentheses.

I also check the alternative measures of some other variables. An alternative sex ratio is to use all males and females aged 20 to 50. It depicts a broader picture of sex imbalance than the sex ratio among 20 to 34 years old men and women. The coefficients are all positive and significant. I use the mean income level to replace the mean log income as the indicator of the average income level in the city. The results for the key explanatory variable are still positive and significant in the first three specifications at the 10% level. Furthermore, using male mean income and the average gender difference in income, the marginal effect is still large and positive, but only significant at the 10% level in the third and fourth specifications.

I further test the baseline model based on different subsamples. To this end, I divide the female sample by age. Respectively, I check the results for women within the age range of 20 to 30 or 20 to 28 in addition to the age range of 20 to 34. I am unable to check the subsample of women aged between 28 and 34 due to a small sample size. When restricted to 20 to 30 years old urban females, the magnitude of the positive effects on female propensity to marry is greater. However, it is not significant for a younger group of women aged from 20 to 28. This implies as the inequality of male income increases, its accelerating effect on marriage is stronger for females between 28 and 30 years old, but not significant for those under the age of 28.

Finally, I test the subsamples separated by education level. Excluding illiterate women, the results for all literate women are still positive and significant at the 5% level for all five specifications. However, after dropping women without a high school diploma, the effect is no longer significant. Hence, the impact of the male income inequality mainly affects women without a high school degree.

3.5.3 Rural Baseline Estimation

The baseline estimation results for rural counties are shown in Table 3.6. All specifications are the same as in the urban estimations, except that there is no province capital fixed effect for rural estimations. In the first specification, the marginal effect of male income inequality is 0.274 significant at the 5% level. In the counties with a higher male income Gini coefficient than the mean (0.437) by the value of 0.01, rural females age 20 to 34 are 0.272 percentage point more likely to be married. Compared with women having a postsecondary degree, women with only compulsory education and lower are more likely to marry before age 34. However, the high school graduates are less likely to marry. Compared with women aged between 28 and 30, younger females are less likely to be married, but elder females especially between 32 to 34 years old are more likely to be married. A higher sex ratio between men and women aged 20 to 50 in a county increases the probability of rural females to be married before 34 years of age.

In the second specification, the marginal effect of male income inequality is 0.648, significant at the 1% level. In the third specification, it is 0.597, significant at the 1% level. In the fourth specification, with the county-specific time trend, the marginal effect becomes 0.831, significant at the 1% level. Finally, in the fifth specification, with both county and year fixed effect as well as the national time trend, the effect of the male income Gini coefficient on female marriage propensity is no longer significant.

The effects of age and education on rural women's marriage propensity are consistent across all five specifications. As women get older in the rural areas, they are more likely to be married. For estimations without year fixed effects, in the same county all literate women have lower propensity to marry than illiterate women. The absolute value of the coefficient is the largest for women with a high school diploma and the smallest for women with only a compulsory education. It indicates that in the same county, high school graduates are less likely to marry compared with women having a postsecondary degree. Women with a postsecondary degree are less likely to marry compared with the women with only a compulsory education. Finally, the illiterate women are the most likely to marry among all women. After controlling for the year fixed effect, we obtain a similar conclusion. That is to say, in a certain year

Table 3.6 Probability of Being "Ever Married" for Rural Women Aged 20-34 (1989-2009)

		(1)	(2)	(3)	(4)	(5)
Dependent variable: 0, never married; 1, ever married.						
Gini of male income		0.272** (0.138)	0.648*** (0.149)	0.597*** (0.145)	0.831*** (0.168)	0.109 (0.128)
Education	Illiteracy	0.176*** (0.038)	–	–	–	0.208*** (0.036)
	Compulsory education	0.119*** (0.030)	-0.113*** (0.036)	-0.115*** (0.036)	-0.126*** (0.037)	0.139*** (0.030)
	High school graduates	-0.087** (0.038)	-0.339*** (0.044)	-0.336*** (0.044)	-0.349*** (0.045)	-0.086** (0.038)
	Postsecondary Degree	–	-0.261*** (0.047)	-0.258*** (0.048)	-0.268*** (0.049)	–
Age	20≤age<22	-0.706*** (0.024)	-.0749*** (0.021)	-0.748*** (0.022)	-0.743*** (0.022)	-0.725*** (0.023)
	22≤age<24	-0.531*** (0.032)	-0.588*** (0.031)	-0.584*** (0.032)	-0.574*** (0.033)	-0.549*** (0.031)
	24≤age<26	-0.343*** (0.036)	-0.396*** (0.039)	-0.390*** (0.039)	-0.373*** (0.040)	-0.358*** (0.036)
	26≤age<28	-0.174*** (0.045)	-0.209*** (0.045)	-0.204*** (0.046)	-0.194*** (0.046)	-0.182*** (0.046)
	28≤age<30	–	-0.037 (0.050)	-0.032 (0.051)	-0.022 (0.051)	–
	30≤age≤32	0.014 (0.050)	–	–	–	0.002 (0.051)
	32≤age≤34	0.133** (0.065)	0.154** (0.062)	0.155** (0.062)	0.165** (0.061)	0.139** (0.063)
Sex ratio (20-50)		0.450*** (0.104)	0.179 (0.136)	0.098 (0.148)	-0.034 (0.155)	0.433*** (0.118)
Mean male log income		0.020 (0.046)	0.115** (0.051)	0.144*** (0.052)	0.177** (0.060)	-0.035 (0.041)
Mean gender difference in log income		0.013 (0.043)	0.007 (0.056)	-0.005 (0.055)	0.038 (0.061)	0.030 (0.045)
Year Dummies		Yes	No	No	No	Yes
Year Trends		National	National	National& Province	National& County	National
Geographic fixed effects		Province	County	County	County	County
Number of county-year clusters		271	271	271	271	271
Number of observations		4,593	4,593	4,593	4,593	4,593

Notes: ***p<0.01, **p<0.05, *p<0.1. Marginal effects of male income inequality on young urban female probability of being ever married are shown. Column (1) to (5) is respectively consistent with the specification (1) to (5) in the baseline regression. Standard errors are corrected by city-year (or community-year for row 2 only) clusters shown in parentheses.

compared to the women with a postsecondary degree, all other women, except for the high school graduates, are more likely to marry especially the illiterate women.

However, the effect of sex ratio on female marriage propensity is only significant after controlling for year fixed effect. Hence, the difference of sex ratio between counties instead of that over time within a county has significant impact on female marriage propensity. The marginal effect of sex ratio on female propensity to marry is 0.45 and significant at the 1% level in the first specification. It is 0.433 and significant at the 1% level in the fifth specification. If the sex ratio becomes bigger than the mean by 0.1, the propensity to marry for women becomes 4.3 percentage points higher than the average level.

The impact of average male income of the county within a year on female marriage propensity is only significant after controlling for county fixed effect and province-specific time trends (or county-specific time trend) in addition to the national time trend. It is not significant after controlling for year fixed effect. Hence, a county should have higher female marriage propensity in the year when the average male income improves compared with the trend level in the county or the trend level in the province. But in the same year, a county with higher average male income compared with other counties does not necessarily have a higher female marriage propensity. The marginal effect of the mean log male income is 0.144, significant at the 1% level in the third specification, and 0.177 significant at the 5% level in the fourth specification. That is to say, a 10% higher value of mean male income within a county compared with the average level in the province leads to a 1.44 percentage point higher propensity to marry, than the average of 62.6 percentage points, for women in the county. While a 10% higher value of mean male income within a county compared with the usual average level in the county leads to 1.77 percentage points higher propensity, than the average probability of 62.9 percentage points, to marry for women in the county. After controlling for the mean values of male log income, the income gap between genders does not have significant impact on female marriage propensity. Implicitly, change in the average female log income does not affect women's marriage propensity.

3.5.4 Rural Robustness Check

Table 3.7 presents the robustness check for the rural sample. Since the rural sample is double the size of the urban sample, the results are generally more robust when compared to the urban sample. There are no changes in the basic results by clustering standard errors in the community-year units instead of the county-year units.

The positive impact of male income inequality on female marriage propensity is more robust with the rural sample. Like in the urban sample, after controlling for the county fixed effect and county-specific time trends, the 50/10 male income ratio has a significant impact on female marriage propensity. The sign becomes positive and as a result that women are more likely to marry when the ratio of median male income over the male income ranked at the bottom 10th percentile becomes higher in the county. Moreover, the coefficients of some other measures for male income inequality become significant only with the rural samples. Unlike the results with urban samples, the standard deviation of male income has a significant positive impact on a rural woman's decision to marry, although the marginal effect is less than half of the magnitude in the baseline results. Furthermore, with county fixed effect and province or county-specific time trends, the 90/10 ratio could also explain the increase of female marriage propensity. In addition, a higher 90/50 ratio among rural men significantly increases a rural woman's propensity to marry. That is to say rural women are more likely to marry as the difference enlarges between male income ranked at the top 10th percentile and the median male income.

Table 3.7 Robustness Check of Probability of Being "Ever Married" for Rural Women Aged 20-34 (1989-2009)

Alternative measures		Marginal effect of male income inequality				
		(1)	(2)	(3)	(4)	(5)
Baseline results of male income Gini		0.272** (0.138)	0.648*** (0.149)	0.597*** (0.145)	0.831*** (0.168)	0.109 (0.128)
Standard errors clustered in community-year		0.272** (0.145)	0.648*** (0.145)	0.597*** (0.148)	0.831*** (0.160)	0.109 (0.152)
Alternative measures of male income inequality	Std of male log income	0.117* (0.068)	0.239*** (0.079)	0.221*** (0.078)	0.273*** (0.089)	0.015 (0.070)
	90/10 ratio	0.000292 (0.000244)	0.000682 (0.000547)	0.0009597* (0.000534)	0.00127** (0.000550)	0.0000361 (0.000344)
	90/50 ratio	0.0388*** (0.0133)	0.0452*** (0.0138)	0.0399*** (0.0134)	0.0539*** (0.0155)	0.02189** (0.0110)
	50/10 ratio	0.000194 (0.000537)	0.00121 (0.00138)	0.00204 (0.00133)	0.00271** (0.00135)	-0.000158 (0.000811)
Sex ratio 20-34		0.231 (0.142)	0.634*** (0.148)	0.579*** (0.144)	0.838*** (0.165)	0.077 (0.131)
Average income	Mean income all 20-50	0.312* (0.189)	0.387* (0.233)	0.457* (0.259)	0.395 (0.271)	0.298 (0.198)
	Mean gender difference in income	0.242 (0.216)	0.419 (0.270)	0.514 (0.286)	0.491* (0.290)	0.236 (0.239)
Restrict samples	Female age 20-30	0.299** (0.151)	0.680*** (0.163)	0.635*** (0.158)	0.909*** (0.181)	0.119 (0.142)
	Female age 20-28	0.341** (0.152)	0.751*** (0.156)	0.716*** (0.150)	0.980*** (0.179)	0.193 (0.136)
	Female with education	0.302* (0.157)	0.699*** (0.158)	0.637*** (0.147)	0.848*** (0.169)	0.183 (0.141)
	Female high school grad	0.255 (0.241)	0.644*** (0.240)	0.595*** (0.227)	0.770*** (0.294)	0.302 (0.261)

Notes: ***p<0.01, **p<0.05, *p<0.1. Marginal effects of male income inequality on young rural female propensity to marry are shown. Column (1) to (5) is respectively consistent with the specification (1) to (5) in the baseline regression. Standard errors are corrected by county-year (or community-year for row 2 only) clusters shown in parentheses.

In comparison with the results for the urban estimation, I find that the coefficients for the sex ratio have exhibited interesting changes in the rural estimation. In the estimation for urban women, the sex ratio for men and women at age 20 to 34 is only significant in the fourth specification with city fixed effect and city-specific time trends. And the sign is negative, which indicates a higher sex ratio among young men and women below 34 deviated from the normal city trend over time is associated with lower marriage propensity for young women. In the robustness check for the urban sample, the measure of sex ratio for men and women aged 20 to 50 always has a positive and significant relationship with female marriage propensity. For rural samples, the relations of sex ratio among men and women below 34 above 20 with female marriage propensity are only significant and positive without year fixed effects. On the contrary, the relationship of sex ratio within a broader age range, say 20 to 50, is only significant and positive with a year fixed effect. Hence, for the rural women's marriage decision, the differences in the sex ratio among young people below 34 over time in the same county matters and the changes in sex ratio for all men and women below 50 across counties in the same year also matters. In both cases, a higher sex ratio is associated with more marriage for rural women.

Finally, results for rural women under different age groups are not the same as urban women. The propensity to marry for rural women below 28 years old also increases when the male income inequality becomes higher in the same county over time, whereas for urban women below 28, there are no significant results. Furthermore, the coefficient of the male wage inequality for 20 to 30 years old women's marriage propensity becomes larger with the rural sample. This is consistent with the notion that rural females tend to marry younger than urban females. Besides, the propensity to marry increases not only for all literate rural women in general, but also for those with the minimum of a high school diploma when the male income inequality enlarges, especially within the same county.

3.6 Discussion

3.6.1 "Reservation Wage" Effect

How does one interpret these results? To begin with, these results indicate that there might be a delay effect from a higher "reservation wage", but the overall results are not dominated by the effect of higher "reservation wage". In general, the results show that as male income inequality increases, female propensity to marry tends to increase in both rural and urban areas. Since a higher reservation wage is mainly due to the increase in the wage inequality at the upper tail of the male wage distribution, the 90/50 ratio would be a better measure to check the impact of a higher "reservation wage" on female marriage propensity (Loughran 2002). According to Table 3.7, the effect of a higher 90/50 ratio is always positive and significant in rural areas, although it is sometimes negative but not significant in urban areas (see Table 3.5). Although the overall impact of male income inequality on urban women's marriage propensity is positive, we also have obtained insignificant results for younger and higher educated urban women with the male income Gini coefficient as the male income inequality measure (see Table 3.5). The assumption is that the increase in "reservation wage" may affect higher educated urban young women more than others. It could be the case that the "reservation wage" effect has mixed with other competing effects, which leads to insignificant results for higher educated urban young women.

If the "reservation wage" effect is not dominant, it could be the case that the needed assumptions for the "reservation wage" effect to be dominant are not met. If the assumption of well-behaved income distribution is violated in this Chinese sample, the "reservation wage" effect is no longer dominant. It is highly possible that the income distribution among males in China is polarized. In this case, the impact of a higher possibility to meet someone above the reservation wage dominates the "reservation wage" impact as discussed in section 3.2.2.

Meanwhile, the assumption that the majority of women are risk neutral may not be true in this sample. Part of the women surveyed had stayed in the same city or county and participated in this survey more than once. They may be more risk adverse than woman migrating out or newly migrating in to a city or county. However, there is no evidence in the sample to verify this point.[28]

3.6.2 Endogeneity

Are the results driven by endogeneity problems due to the lack of control on the female economic status? Theoretically, the link between male income inequality and female marriage formation can be formed by a missing variable regarding the women's economic performance. On one hand, women's career development will affect their marital timing. On the other hand, women's wage changes can affect male wage inequality (Acemoglu, Autor, and Lyle 2004). Female wage gains could be male wage losses, and female wage distribution is linked to male wage distribution when the overall wage distribution remains stable over time (Fortin and Lemieux 2000). Besides, women's earning power determines their husbands' expected future earnings (Sweeney and Cancian 2004).

To check this endogeneity problem, in Table 3.8 I add the individual woman's income or log income control variable in addition to the female age and her education level into the baseline regression. Due to missing values from women without work, the number of observations drops from 2,241 in the baseline to 1,697 by using income and 1,671 by using log income. With both measures, the magnitude for the coefficient of male income Gini coefficient is positive and only a little bit smaller than that in the baseline results for the urban sample, but greater for rural sample. It is possible that the effect of male income inequality on rural female marriage propensity is underestimated due to the lack of control for women's income. There is little evidence, however, to indicate that the observed effects of male income inequality are mainly due to the impact of the female own income on her marriage decision.

3.6.3 Reverse Causality

The causality could run the opposite way, which would be that the change in income inequality among men is due to the change in the propensity to marry among women. For example, consider that the marriage rate among young men becomes smaller because women choose to marry later. It could be the case if women prefer to marry men with similar or older age (Buss 1989; Higgins et al. 2002). Meanwhile, there is wage premium for men after they marry (Korenman and Neumar 1991; Gray 1997; Chun and Lee 2001; Ahituv and Lerman 2007). We also know that there is larger variation in wage for single men compared with married men (Blackburn 1990). Hence, less marriage among young men can be associated with changes in the inequality levels of male earnings.

In order to check this reverse causality issue, I only use the Gini coefficient among married males aged 20 to 50 as an explanatory variable (see Table 3.8). Usually, it is easier for a woman to observe or detect the change in the income inequality of all men as a whole than that of young men only. Among all 20,457 male observations with income between age 20 and 50, single men

[28] There are no notable differences of the income, education level, and parents' cadres status (whether they are public servants or not) between women who stay in the sample and those who are either newly added or only occurring once in the survey.

are only 15.39%. The income inequality of married men can substantially change female marriage decision-making in the same way as the overall male income inequality does. Using married men's income Gini coefficient, results obtained are very similar with the baseline results. Most of coefficients for both urban and rural female samples have greater magnitudes compared with those obtained from the baseline estimations. Theoretically, female marriage propensity cannot affect the male income inequality among married men. Therefore, change in female marital propensity is the result, not the cause of male income inequality.

Table 3.8 Endogeneity, Reverse Causality and Other Mechanisms

Key explanatory variable		Marginal effect of the key explanatory variable on female being "ever married" (age 20-34)				
		(1)	(2)	(3)	(4)	(5)
Urban Female 20-34						
Male income Gini	Control female log income	0.289 (0.212)	0.592** (0.268)	0.544* (0.297)	0.598* (0.334)	0.337 (0.226)
	Control female income	0.270 (0.205)	0.576** (0.260)	0.557* (0.290)	0.625* (0.327)	0.302 (0.219)
Married male income Gini		0.385* (0.214)	0.642*** (0.230)	0.707*** (0.255)	0.725*** (0.251)	0.376** (0.194)
Income Gini for all people (men & women)		0.139 (0.199)	0.448 (0.291)	0.543* (0.293)	0.580* (0.298)	0.115 (0.215)
Female income Gini		-0.159 (0.137)	0.104 (0.223)	0.150 (0.213)	0.253 (0.213)	-0.184 (0.145)
Rural Female 20-34						
Male income Gini	Control female log income	0.352** (0.158)	0.750*** (0.163)	0.700*** (0.160)	0.929*** (0.178)	0.223 (0.143)
	Control female income	0.321** (0.157)	0.749*** (0.166)	0.686*** (0.157)	0.899*** (0.177)	0.194 (0.143)
Married male income Gini		0.307** (0.140)	0.670** (0.144)	0.633*** (0.137)	0.797*** (0.160)	0.187 (0.123)
Income Gini for all people (men & women)		0.323* (0.180)	0.894*** (0.197)	0.856*** (0.186)	1.084*** (0.222)	0.162 (0.157)
Female income Gini		0.063 (0.170)	0.474** (0.222)	0.453** (0.216)	0.598** (0.242)	0.088 (0.164)

Notes: ***p<0.01, **p<0.05, *p<0.1. Marginal effect of different income inequality measurements on age 20 to 34 urban/rural female's marriage propensity is shown. Column (1)-(5) is respectively consistent with the specification one to five in the baseline regression. Controls are the same with baseline estimations if change is not mentioned. Standard errors are in parentheses either corrected by city-year or county-year clusters.

3.6.4 Risk and Uncertainty

Could the link between male income inequality and female marriage propensity be formed due to their association with the income risk and social uncertainty? We know that the income from men is usually the most important income source for a Chinese family. An enlarged income inequality among men would increase the income risk and social uncertainty for the society. It is very likely that enlarged income inequality of males just reflects part of the exaggeration of overall income inequality in the society. Therefore, income inequality affects female marriage formation through the impact of income risk and social uncertainty on marriage, the income inequality among men and women should both change the propensity for females to marry.

In order to test this hypothesis, I replace the Gini coefficient of male income with that of all individuals aged from 20 to 50. The results are relatively weaker for urban residents, but stronger for rural residents. For specifications (3) and (4), controlling for city fixed effect and province or city specific time trend, the total income inequality has positive effect on marriage propensity of young females, which is significant at the 10% level. For rural females, an increase in the overall income Gini coefficient in a county within a year could lead to an increase in the propensity to marry, which is better than the male income Gini coefficient under all five specifications.

Likewise, if the income risk and social uncertainty is the mechanism for male income inequality to impact female marriage formation, female income inequality should have a similar effect on female marriage propensity as male income inequality does. I calculate female income inequality using women aged from 20 to 50. In urban areas, female income inequality has no effect on their marriage propensity. But in rural areas, the increase in the female income inequality also significantly increases young females' propensity to marry.

In summary, a rural woman's marriage propensity can be affected by the increase in income inequality among only males, only females and all people. By contrast, the urban female marital decision-making is only affected by the male income inequality and the income inequality among all people. Given all that, income risk and social uncertainty is considered to be a more influential factor for rural women than for urban women in linking their marriage propensity to the income inequality.

3.6.5 Men's Marital Decision

Men's marital decision cannot be neglected. Women's economic gains are starting to resemble those of men in a modern family (Oppenheimer 1988; Okun 2001; Sweeney 2002). Men may follow the same strategy as women do in a marital search. A marriage trend may result from similar decision-making processes used by both men and women. Other than searching, it usually takes time for high ability men to signal their quality to prospective women in order to marry a desired partner (Bergstrom and Bagnoli 1993). Some men prefer to spend more time in their career than in maintaining a relationship. While, others may prefer to marry early, since marriage also improves the earnings of men.

As shown in Table 3.9, I regress a young man's marriage propensity on the income inequality among men, women and all people. I include men aged between 22 and 34 years old.[29] The results show that the urban/rural men's marriage propensity increases when the income inequality among urban/rural men increases except for the last specification where I control for both city (county) and year fixed effects and the city-specific (county-specific) time trend. It is true even after controlling for the log value of the male income, given there is no year fixed effect. These results are not due to reverse causality since it is still robust after using the

[29] For urban estimation, the samples size is 1,953 and 1,540 after controlling own income. For rural estimation, it becomes 4,049 and 3,033 after controlling own income.

Table 3.9 Probability of Being "Ever Married" for Urban or Rural Men Aged 22-34 (1989-2009)

Key explanatory variable		Marginal effect on male being "ever married" (1)	(2)	(3)	(4)	(5)
Urban men						
Male income Gini	Baseline	0.376* (0.224)	0.505* (0.260)	0.607** (0.273)	0.515* (0.284)	0.320 (0.209)
	Control log income	0.382 (0.240)	0.475* (0.280)	0.579* (0.305)	0.665** (0.340)	0.255 (0.232)
Married male income Gini		0.367* (0.197)	0.479** (0.230)	0.562** (0.241)	0.502* (0.263)	0.257 (0.190)
Income Gini for all people (men & women)		0.343 (0.253)	0.455 (0.301)	0.610** (0.269)	0.539* (0.326)	0.208 (0.227)
Female income Gini		0.038 (0.218)	0.187 (0.234)	0.253 (0.226)	0.278 (0.240)	-0.073 (0.174)
Rural men						
Male income Gini	Baseline	0.292* (0.162)	0.671*** (0.194)	0.630*** (0.142)	0.854*** (0.150)	0.189 (0.144)
	Control log income	0.243 (0.159)	0.753*** (0.164)	0.718*** (0.152)	0.926*** (0.174)	0.117 (0.141)
Married male income Gini		0.242 (0.148)	0.616*** (0.143)	0.587*** (0.140)	0.787*** (0.147)	0.165 (0.130)
Income Gini for all people (men & women)		0.312 (0.203)	0.880*** (0.207)	0.855*** (0.189)	1.140*** (0.216)	0.223 (0.180)
Female income Gini		0.076 (0.187)	0.417 (0.226)	0.421* (0.219)	0.611* (0.249)	0.082 (0.177)

Notes: ***p<0.01, **p<0.05, *p<0.1. Marginal effects of different income inequality measurements on age 22 to 34 urban or rural male marital decision-making are shown. Column (1) to (5) is respectively consistent with the specification (1) to (5) in the baseline regression. Controls are individual male education and age and sex ratio, mean log income of males and average log income gender difference in marriage market for the baseline. Standard errors are in parentheses either corrected by city-year or county-year clusters.

Gini coefficient of income among only married men. The Gini coefficient for income among all people also has a positive effect on the marital propensity of men in both rural and urban samples after controlling for a city/county fixed effect and the province-specific or city-specific (county-specific) time trends. However, only in the rural sample, after controlling for the county fixed effect and the province-specific or county-specific time trend, men will have higher marriage propensity when there is higher income inequality among women.

The reasons for these results from male samples might be the following. Firstly, men follow the same marital search strategy as women do especially in rural areas. As income inequality of women increases, there are competing effects on the possibility to marry. In rural areas, the promotion effect caused by the higher possibility for a man to accept an offer at a given "reservation wage" is dominant. Moreover, men just want to marry and get more secure as women do when inequality is higher in the society. Therefore, when income inequality of men becomes larger, both rural and urban young men have higher demand for marriage. Finally, more men can get married because more women decide to marry. The impact of the male income inequality or the income inequality among all people on the urban/rural women leads to the change in the marriage propensity observed for urban/rural men.

3.7 Conclusion

This paper shows that Chinese women aged between 20 and 34 are more likely to marry, when income inequality among men is higher in a city or county in a year, so called a "marriage market". The results are obtained by applying a linear probit model to around 2,200 female observations from 18 urban cities and 4,600 female observations from 36 rural counties over year 1989, 1991, 1993, 1997, 2000, 2004, 2006 and 2009. Results are conditional on the woman's age, education and her own income as well as the sex ratio, mean male income level, and average income gap between genders in the marriage market. The effects are usually stronger after controlling for the province-specific time trend or the city/county-specific time trend. A one standard deviation increase in the Gini coefficient for men in the county or city increases the propensity to marry for women aged from 20 to 34 by 5.8 percentage points in urban areas and by 6.9 percentage points in rural areas (see Table 3.4 specification (4)).

Two theories could explain the positive effect of male income inequality on female marriage propensity. The first is based on the marital search theory given a more polarized income distribution in China. The effect of higher probability to meet someone with income above the "reservation wage" level dominates the delay effect of a higher "reservation wage" when there is higher male income inequality. Another reason is that an increase in the income risk and social uncertainty due to a higher income inequality makes marriage more attractive for both young Chinese men and women.

The link between male income inequality and female higher probability to marry is not due to the endogeneity from the woman's own income. It is not caused by reverse causality. The results are also robust with respect to a variety of checks with different variable measurements and subsamples.

In general, the effects are more significant for rural, lower educated, and older women. Results for rural areas are not sensitive to the selection of age range and education levels. However, in urban areas, women with education higher than that of a high school graduate degree and those before age 28 do not have a significant increase in their marriage propensity under a higher income inequality of men. In rural area, a higher 90/50 ratio increases female marriage propensity, whereas a higher 50/10 ratio reduces marriage propensity. For urban females, when making marital decisions, the change in income inequality among all individuals or only among women is not as important as the change in the income inequality among men. But a rural woman would change her marriage propensity based on a different level of income inequality among all men, or among all women, or even among all people.

Male marriage propensity may interact with female marriage propensity. Young men before age 34 would like to accelerate marriage under higher income inequality among men or among all people, which is not because of the reverse causality. Higher female income inequality also increases a rural man's propensity to marry.

In conclusion, this paper demonstrates the impact of income inequality on the marriage formation in China. It provides a new perspective on understanding the influences of income inequality and the reasons of marriage propensity changes in China. It emphasizes the importance of family functions and the higher desirability of marriage under severely rising income inequality. Given a unique socioeconomic background in China, marriage is more desirable under a higher income inequality, especially for the rural residences, the elder and lower educated men or women. Though the propensity to marry is higher under substantial increase of income inequality, the actual marriage rate is not necessarily higher given the effects of some other factors such as an unbalanced sex ratio. For rural men with low income the high sex ratio decreases their marriage rate. Therefore, the rising income inequality adds more conflicts in the reality of marriage formation in China.

CHAPTER 4. CONCLUSION

This dissertation builds upon the work of Loughran (2002) and the Gould and Paserman (2003), which associates female marital decision-making with male wage inequality in the U.S. It has been argued that under a marital search framework, women will delay marriage because the minimum acceptable wage of their potential partners for a woman to marry becomes higher in the situation that male wage inequality becomes larger in a metropolitan area in a year in the U.S.

However, to claim these effects to be causal, there are still endogeneity and reverse causality problems. Previous studies (Loughran 2002; Gould and Paserman 2003; Kuo 2008) only address these problems by controlling for the metropolitan area fixed effect and the time trend in the county or in each metropolitan area. To solve this problem, in Chapter 2, I apply skill-biased technological shock in a metropolitan area in a year as an instrument for the male wage inequality following the example of Mocan and Unel (2011). To facilitate the usage of this instrument, I measure the male wage inequality by the wage gap between high and low educated men. High educated is defined as with some college or higher education and vice versa.

As a result, I discover that a low educated woman's marriage propensity becomes lower but a high educated woman's marriage propensity becomes higher when there is an increase in the wage ratio between high and low educated men. The data is from the 1990 and 2000 U.S. Censuses as well as the 2007 American Community Survey. This helps to explain why the marriage decline for high educated women becomes slower than that for low educated women. The impact of the male wage gap between high and low educated men on marriage propensity of women roughly accounts for 1/5 of the changes in the marriage rate for women in the last two decades.

In addition, in the chapter 3, I apply similar models as in Gould and Paserman (2003) to Chinese data in order to find more evidence about the underlying impact of the male income inequality on female marriage propensity. It turns out that there is an opposite impact in China by utilizing data from the Chinese Health and Nutrition Survey. Instead of a marital delay, I find that for women aged between 20 and 34 a one-standard-deviation increase in the Gini coefficient of male income increases a woman's probability of being married by 5.8 percentage points in urban areas and 6.9 percentage points in rural areas from 1989 to 2009. This is conditional on the female personal characteristics such as the woman's age and education and the city or county characteristics including the sex ratio, overall income level and gender wage gap. In addition, I control for the year fixed effect and/or city (county) fixed effect as well as different levels of regional-specific year trends. The results could be explained by the polarized income distribution and the essential functions of marriage as a social security net-work and a risk sharing union in the Chinese society, which are not consistent with the U.S. context.

To summarize, changes in the earning inequality among men could significant impact women's decision to get married. However, the impact exhibits different features in a western country like the U.S. and in an eastern country like China. In the U.S., the groups of people who are relatively disadvantaged in the labor market, such as those with less skill or lower education, seem to become less likely to marry under a larger wage gap between high and low skilled workers. The skill-biased technological change not only drives up the wage gap between high and low educated men but also close the original gap between high and low educated women's marriage rates. White and non-white women are both affected by the wage gap between high and low educated men, where the impact seems to be larger for non-white women.

In China, most women are more likely to marry in response to a higher male income inequality. Both rural and urban Chinese women react to the change in the male income inequality especially to the variation of the male income inequality over time within a rural county or an urban city. And the impact is more robust for rural women. Polarization in the income distribution in the income inequality further boosts the demand for marriage. Not only the male income inequality but also the income inequality among all men and women could increase people's propensity to marry in the rural area or for people without a high school degree in the urban area.

An explanation of the opposite impact found from the two countries could rely on the differences in the assumptions of the marital search theory. Specifically, when there are changes in the shape of the male wage/income distribution, the impact of the male wage/income inequality on female marriage propensity in the marital search process could be different according to the search theory. Meanwhile, the development of social security system interacts with the impact of income or wage inequality on marriage formation. Countries with a better social security system like the U.S., when compared with China, may experience more severe marriage delay.

REFERENCES

Acemoglu, Daron. 1998. “Why Do New Technologies Complement Skills? Directed Technical Change and Wage Inequality. " *The Quarterly Journal of Economics* 113(4):1055-89.

———. 2011. “Skills, Tasks and Technologies: Implications for Employment and Earnings.” *Handbook of Labor Economics* 4:1043-1171.

Acemoglu, Daron, and David Lyle. 2004. "Women, War, and Wages: The Effect of Female Labor Supply on the Wage Structure at Midcentury." *Journal of Political Economy* 112(3): 497-551.

Ahituv, Avner, and Robert I. Lerman. 2007. “How Do Marital Status, Work Effort, and Wage Rates Interact?” *Demography* 44(3):623-647.

Akerlof, George A., Janet L. Yellen, and Michael L. Katz. 1996. "An Analysis of Out-of-Wedlock Childbearing in the United States." *The Quarterly Journal of Economics* 111(2):277-317.

Angrist, Josh. 2002. "How Do Sex Ratios Affect Marriage and Labor Markets? Evidence from America's Second Generation." *The Quarterly Journal of Economics* 117(3):997-1038.

Autor, David H., Lawrence F. Katz, and Melissa S. Kearney. 2008. "Trends in US Wage Inequality: Revising the Revisionists." *The Review of Economics and Statistics* 90(2):300-23.

Becker, Gary S. 1973. "A Theory of Marriage: Part I." *The Journal of Political Economy* 81(4):813-46.

Becker, Gary Stanley, and Gary S. Becker. 2009. *A Treatise on the Family*. Cambridge, Mass.: Harvard University Press.

Bergstrom, Theodore C., and Mark Bagnoli. 1993. "Courtship as a Waiting Game." *Journal of Political Economy* 101(1):185-202.

Bian, Yanjie. 2002. "Chinese Social Stratification and Social Mobility." *Annual Review of Sociology* 28(1):91-116.

Blackburn, McKinley L. 1990. "What Can Explain the Increase in Earnings Inequality Among Males?" *Industrial Relations: A Journal of Economy and Society* 29(3):441-56.

Browning, Marin, Pierre-Andre Chiappori, and Yoram Weiss. 2011. "*Family Economics*." Tel Aviv University, Unpublished Textbook Manuscript.

Buchinsky, Moshe, and Jennifer Hunt. 1999. "Wage Mobility in the United States." *Review of Economics and Statistics* 81(3):351-68.

Burdett, Kenneth, and Jan I. Ondrich. 1985. "How Changes in Labor Demand Affect Unemployed Workers." *Journal of Labor Economics* 3(1):1-10.

Buss, David M. 1989. "Sex Differences in Human Mate Preferences: Evolutionary Hypotheses Tested in 37 Cultures." *Behavioral and Brain Sciences* 12(1):1-49.

Card, David, and John E. DiNardo. 2002. "Skill Biased Technological Change and Rising Wage Inequality: Some Problems and Puzzles." *Journal of Labor Economics* 20(4):733-82

Chen, Xi, and Xiaobo Zhang. 2009. "The Distribution of Income and Well-being in Rural China: A Survey of Panel Data Sets, Studies and New Directions." Munich Personal RePEc Archive Paper No. 20587.

Chiappori, Pierre-Andre, Murat Iyigun, and Yoram Weiss. 2009. "Investment in Schooling and the Marriage Market." *The American Economic Review* 99(5):1689-1713.

Choe, Jongmook. 2008. "Income Inequality and Crime in the United States." *Economics Letters* 101(1):31-3.

Choo, Eugene, and Aloysius Siow. 2006. "Who Marries Whom and Why." *Journal of Political Economy* 114(1):175-201.

Chun, Hyunbae, and Injae Lee. 2001. "Why Do Married Men Earn More: Productivity or Marriage Selection?" *Economic Inquiry* 39(2):307-19.

Coale, Ansley J. 1992. "Age of Entry into Marriage and the Date of the Initiation of Voluntary birth control." *Demography* 29(3):333-41.

Coughlin, Tristan P., and Scott D. Drewianka. 2011. "Can Rising Inequality Explain Aggregate Trends in Marriage? Evidence from US States, 1977-2005." *The BE Journal of Economic Analysis & Policy* 11(1):1935-1682.

Dasgupta, Monica, Avraham Ebenstein, and Ethan J. Sharygin. 2011. "China's Bride Shortage & Upcoming Challenges for Elderly Men." Development Research Group. The World Bank.

DiNardo, John, Nicole M. Fortin, and Thomas Lemieux. 1996. "Labor Market Institutions and the Distribution of Wages, 1973-1992: A Semiparametric Approach." *Econometrica* 64(5):1001-44.

Fortin, Nicole M., and Thomas Lemieux. 2000. "Are Women's Wage Gains Men's Losses? A Distributional test." *The American Economic Review* 90 (2000):456-60.

Geweke, John, and Michael Keane. 2000. "An Empirical Analysis of Earnings Dynamics Among Men in the PSID: 1968–1989." *Journal of Econometrics* 96(2):293-356.

Goldin, Claudia, and Lawrence F. Katz. 2000. The Power of the Pill: Oral Contraceptives and Women's Career and Marriage Decisions. Working Paper No. 7527. Cambridge, Mass.: National Bureau of Economic Research.

Goldstein, Joshua R., and Catherine T. Kenney. 2001. "Marriage Delayed or Marriage Forgone? New Cohort Forecasts of First Marriage for US Women." *American Sociological Review* 66(4):506-19.

Gottschalk, Peter, Robert Moffitt, Lawrence F. Katz, and William T. Dickens. 1994. "The Growth of Earnings Instability in the US Labor Market." *Brookings Papers on Economic Activity* 1994(2):217-72.

Gottschalk, Peter. 1997. "Inequality, Income Growth, and Mobility: The Basic Facts." *The Journal of Economic Perspectives* 11(2):21-40.

Gould, Eric D., and M. Daniele Paserman. 2003. "Waiting for Mr. Right: Rising Inequality and Declining Marriage Rates." *Journal of Urban Economics* 53(2):257-81.

Gray, Jeffrey S. 1997. "The Fall in Men's Return to Marriage: Declining Productivity Effects or Changing Selection?" *Journal of Human Resources* 32(3):481-504.

Greenstone, Michael, and Adam Looney. 2012. "The Marriage Gap: The Impact of Economic and Technological Change on Marriage Rates." *Brookings Institution, February 3*.

Gustafsson, Björn, and Li Shi. 2002. "Income Inequality Within and Across Counties in Rural China 1988 and 1995." *Journal of Development Economics* 69(1):179-204.

Han, Hongyun. 2010. "Trends in Educational Assortative Marriage in China from 1970 to 2000." *Demographic Research* 22(24):733-70.

Hess, Gregory D. 2004. "Marriage and Consumption Insurance: What's Love Got to Do with It?" *Journal of Political Economy* 112(2):290-318.

Higgins, Louise T., Mo Zheng, Yali Liu, and Chun Hui Sun. 2002. "Attitudes to Marriage and Sexual Behaviors: A Survey of Gender and Culture Differences in China and United Kingdom." *Sex Roles* 46(3-4):75-89.

Hogan, Dennis P. 1978. "The Effects of Demographic Factors, Family Background, and Early Job Achievement on Age at Marriage." *Demography* 15(2):161-75.

Holden, Karen C., and Pamela J. Smock. 1991. "The Economic Costs of Marital Dissolution: Why Do Women Bear a Disproportionate cost?" *Annual Review of Sociology* 17(1): 51-78.

Keeley, Michael C. 1977. "The Economics of Family Formation." *Economic Inquiry* 15(2):238-50.

Korenman, Sanders, and David Neumark. 1991. "Does Marriage Really Make Men More Productive?" *Journal of Human Resources* 26(2):282-307.

Kuo, Yu-Chen. 2008. "Wage Inequality and Propensity to Marry after 1980 in Taiwan." *International Journal of Business and Economics* 7(3):231-48.

Leamer, Edward E. 1996. "Wage Inequality from International Competition and Technological Change: Theory and Country Experience." *The American Economic Review* 86(2):309-14.

Lewis, Susan K., and Valerie K. Oppenheimer. 2000. "Educational Assortative Mating Across Marriage Markets: Nonhispanic Whites in the United States." *Demography* 37(1): 29-40.

Li, R. 1985. "An Analysis of Age At First Marriage of China's Population." *Population research (Peking, China)* 2(4):27-35.

Li, Hongbin, and Yi Zhu. 2006. "Income, Income Inequality, and Health: Evidence from China." *Journal of Comparative Economics* 34(4):668-93.

Li, Peilin. 1996. "Has China Become Polarized?" *Chinese Economy* 29 (3):73-6.

Lichter, Daniel T., Felicia B. LeClere, and Diane K. McLaughlin. 1991. "Local Marriage Markets and the Marital Behavior of Black and White Women." *American Journal of Sociology* 96(4): 843-67.

Lichter, Daniel T., Diane K. McLaughlin, George Kephart, and David J. Landry. 1992. "Race and the Retreat from Marriage: A Shortage of Marriageable Men?" *American Sociological Review* 57(6):781-99.

Loughran, David S. 2002. "The Effect of Male Wage Inequality on Female Age at First Marriage." *Review of Economics and Statistics* 84(2):237-50.

Mare, Robert D. 1991. "Five Decades of Educational Assortative Mating." *American Sociological Review* 56(1):15-32.

Maurer-Fazio, Margaret, James W. Hughes, and Dandan Zhang. 2005. "Economic Reform and Changing Patterns of Labor Force Participation in Urban and Rural China." Working Paper No. 787. Ann Arbor, MI.: The William Davidson Institute at the University of Michigan

Meng, Lei. 2009. "Bride Drain: Rising Female Migration and Declining Marriage Rates in Rural China." Research Paper, Xiamen University Department of Economics.

Meng, Xin. 2004. "Economic Restructuring and Income Inequality in Urban China." *Review of Income and Wealth* 50(3):357-79.

Mocan, Naci H., and Bulent Unel. "Skill-biased Technological Change, Earnings of Unskilled Workers, and Crime." Working Paper No. 17605. Cambridge, Mass.: National Bureau of Economic Research.

Moffitt, Robert A., and Peter Gottschalk. 2011. Trends in the Transitory Variance of Male Earnings in the US, 1970-2004. Working Paper No. 16833. Cambridge, Mass.: National Bureau of Economic Research.

Montgomery, Mark R., Paul PL Cheung, and Donna B. Sulak. 1988. "Rates of Courtship and First Marriage in Thailand." *Population Studies* 42(3):375-88.

Mortensen, Dale T. 1986. "Job Search and Labor Market Analysis." *Handbook of Labor Economics* 2:849-919.

Mu, Zheng, and Yu Xie. Marital. 2014. "Marital Age Homogamy in China: A Reversal of Trend in the Reform Era?" *Social Science Research* 44:141-57.

Muhsam, H. V. 1974. "The Marriage Squeeze." *Demography* 11(2):291-99.

Okun, Barbara S. 2001. "The Effects of Ethnicity and Educational Attainment on Jewish Marriage Patterns: Changes in Israel, 1957–1995." *Population Studies* 55(1):49-64.

Oppenheimer, Valerie Kincade, Matthijs Kalmijn, and Nelson Lim. 1997. "Men’s Career Development and Marriage Timing during a Period of Rising Inequality." *Demography* 34(3):311-30.

Oppenheimer, Valerie Kincade. 1988. "A Theory of Marriage Timing." *American Journal of Sociology* 94(3):563-91.

———. 1994. "Women's Rising Employment and the Future of the Family in Industrial Societies." *Population and Development Review* 20(2):293-342.

———. 1997. "Women's Employment and the Gain to Marriage: The Specialization and Trading Model." *Annual Review of Sociology* 23:431-453.

———. 2003. "Cohabiting and Marriage During Young Men’s Career-Development Process." *Demography* 40(1):127-149.

Pencavel, John. 1998. "Assortative Mating by Schooling and the Work Behavior of Wives and Husbands." *The American Economic Review* 88(2):326-329.

Perkins, Dwight H. 2011. "History, Politics and 30 Years of Development and Reform." *Reform and Development in China: What Can China Offer the Developing World*. Ed. Yang Yao, and Ho-Mou Wu. Routledge, 45-70.

Qin, Duo, Marie Anne Cagas, Geoffrey Ducanes, Xinhua He, Rui Liu, and Shiguo Liu. 2009. "Effects of Income Inequality on China's Economic Growth." *Journal of Policy Modeling* 31(1):69-86.

Raymo, James M. 2003. "Educational Attainment and the Transition to First Marriage Among Japanese Women." *Demography* 40(1):83-103.

Ru, Xin, Xueyi Lu and Peilin Li. 2007. *2008: Analysis and Forecast of Chinese Society*. Beijing, China: Social Sciences Academic Press.

Santow, Gigi, and Michael Bracher. 1994. "Change and Continuity in the Formation of First Marital Unions in Australia." *Population Studies* 48(3):475-496.

Schoen, Robert. 1983. "Measuring the Tightness of a Marriage Squeeze." *Demography* 20(1):61-78.

Seitz, Shannon. 2009. "Accounting for Racial Differences in Marriage and Employment." *Journal of Labor Economics* 27(3):385-437.

Seitz, Shannon, Aloysius Siow, and Eugene Choo. 2010. "The Collective Marriage Matching Model: Identification, Estimation and Testing." Meeting Paper No. 267. Society for Economic Dynamics.

Siow, Aloysius. 2009. "Testing Becker's Theory of Positive Assortative Matching." Working Paper No.356. University of Toronto

Santow, Gigi, and Michael Bracher. 1994. "Change and Continuity in the Formation of First Marital Unions in Australia." *Population Studies* 48(3):475-496.

Schmidt, Lucie. 2008. "Risk Preferences and the Timing of Marriage and Childbearing." *Demography* 45(2):439-460.

Schoen, Robert. 1983. "Measuring the Tightness of a Marriage Squeeze." *Demography* 20(1):61-78.

Smock, Pamela J., and Wendy D. Manning. 1997. "Cohabiting Partners' Economic Circumstances and Marriage." *Demography* 34(3):331-341.

Smyth, Russell, and Xiaolei Qian. 2008. "Inequality and Happiness in Urban China." *Economics Bulletin* 4(23):1-10.

Stevenson, Betsey, and Justin Wolfers. 2007. "Marriage and Divorce: Changes and Their Driving Forces." Working Paper No.12944. Cambridge, Mass.: National Bureau of Economic Research.

Sweeney, Megan M., and Maria Cancian. 2004. "The Changing Importance of White Women's Economic Prospects for Assortative Mating." *Journal of Marriage and Family* 66(4):1015-28.

Thorbecke, Erik, and Chutatong Charumilind. 2002. "Economic Inequality and Its Socioeconomic Impact." *World Development* 30(9):1477-95.

Unel, Bulent. 2010. "Analyzing Skilled and Unskilled Labor Efficiencies in the US." *Journal of Macroeconomics* 32(4):957-67.

Watson, Tara, and Sara McLanahan. 2011. "Marriage Meets the Joneses Relative Income, Identity, and Marital Status." *Journal of Human Resources* 46(3):482-517.

Wilson, William Julius. 2012. *The Truly Disadvantaged: The Inner City, the Underclass, and Public Policy*. University of Chicago Press.

Xia, Yan R., and Zhi G. Zhou. 2003. "The Transition of Courtship, Mate Selection and Marriage in China." *Mate Selection Across Cultures* 231-246.

Xie, Yu, James M. Raymo, Kimberl Goyette, and Arland Thornton. 2003. "Economic Potential and Entry into Marriage and Cohabitation." *Demography* 40(2):351-367.

Xu, Lixin, Christine Qiang, and Limin Wang. 2002 "The Timing of Marriage in China." Available at SSRN 368402. http://dx.doi.org/10.2139/ssrn.368402.

Yang, Dennis Tao. 1998. "Urban-biased Policies and Rising Income Inequality in China." *The American Economic Review* 89(2):306-10.

Ye, Wenzhen. 1992. "China's ‘Later’Marriage Policy and Its Demographic Consequences." *Population Research and Policy Review* 11(1):51-71.

Zhang, Xiaobo, and Ravi Kanbur. 2001. "What Difference Do Polarisation Measures Make? An Application to China." *Journal of Development Studies* 37(3):85-98.

Zhao, Zhongwei. 2006. "Income Inequality, Unequal Health Care Access, and Mortality in China." *Population and Development Review* 32(3):461-483.

Zhu, Wei Xing, Li Lu, and Therese Hesketh. 2009. "China’s Excess Males, Sex Selective Abortion, and One Child Policy: Analysis of Data from 2005 National Intercensus Survey." *BMJ: British Medical Journal* 338(7700):920-23.

APPENDIX

A. Supplementary Materials for Chapter 2 Section 2.4 Data

I keep individuals aged between 17 to 65 years old in the sample. The years of education are 0 (N/A or no schooling), 2 (nursery school to grade 4), 6 (grade 5,6,7 and 8), 9 (grade 9), 10 (grade 10), 11 (grade 11), 12 (grade 12), 13 (1 year of college), 14 (2 years of college), 15 (3 years of college), 16 (4 years of college) and 18 (5+ years of college). "Years of experience" is the minimum value between possible years of experience (age minus 16 minus years of education) and legal years of working (age minus 16). I drop these with negative value of experiences[30]. In addition to original race indicator, I also categorize all other races except for white into one big race category called "non-white". There are occupation dummies for 7 aggregate categories based on 1990 classification method. They are respectively, "Managerial and Professional", "Technical, Sales, and Administrative", "Service", "Farming, Forestry, and Fishing", "Precision Production, Craft, and Repairers", "Operatives and Laborers", and "Non-occupational responses".

In order to calculate the wage gap between high and low educated men, I divide samples into 32 efficiency groups, which were jointly defined by 2 genders, 4 education levels and 4 experience groups. The four education category indicators are respectively less than or equal to 11 years, 12 years, 13 to 15 years, and equal to or more than 16 years of education. The four experience groups are 0 to 9, 10 to 19, 20 to 29 and 30 or more years of experience[31]. High educated (skilled) individuals are defined as persons with some college and more education, i.e. education categories 3 and 4, and low educated (skilled) are defined as persons with high school or lower education, i.e. education categories 1 and 2.

Further, for the later calculation of race-specific male wage gap, I will also add the race indicator as another dimension. There are two broad race categories, white and non-white. With this additional dimension on race, I create 64 efficiency groups. For instance, group 1 contains white male with less than or equal to 11 years of education and 0 to 9 years of experience and group 2 is non-white male with same other features as group 1. Although, experience group indicators are not used explicitly in the calculation, this division ensures the precision of the constructed variables calculated within each group, which follows related literatures, such as Autor, Katz and Kearney (2008).

For simplicity, let me call the current sample "raw sample". The general version of sex ratio is calculated by the number of men aged 21 to 40 divided by the number of women aged 21 to 35 in "raw sample". I also separate samples by race and education groups and calculate the corresponding within group sex ratio. For example, for the estimation of white female with higher education, I calculate the sex ratio between high educated white men aged 21 to 40 and high educated white women aged 21 to 35.

[30] It is possible to have -1 or -2 experience value due to slightly different beginning age of education or the period of education for a certain degree level. I keep these observations and give them years of experience value of 0.

[31] To be specific, the first 16 groups are for males and other 16 groups are for females. Group 1-8 and 17-24 are low educated persons who has fewer than or equal to 12 years of education, equivalent to high school graduates. Group 9-16 and 25-32 are high educated persons, who have at least some college education. Group 1,9,17 and 25 are within the first experience group, 2,10,18 and 26 are within the second experience group and so on.

I adopt workers worked for full hours and full time (FHFT) last year, whose weekly hours worked are more than 35 hours and weeks worked are more than 40 weeks. I exclude self-employed individuals and those with no salary and income wage. I calculate hourly wages by dividing the total wage and salary income by the product of usual hours worked each week and weeks worked last year. According to Autor, Katz and Kearney (2008), I "drop the bottom 1% of hourly earners and multiply hourly wage of top-coded earners by 1.5". I "limit the maximum hourly wage to 1.5 times the maximum annual income amount divided by 1750 (35 hours per week for 50 weeks per year)". The top-coding values are \$140,000 in the 1990 Census, \$175,000 in the 2000 Census and \$200,000 in the 2007 ACS[32]. I adjust the nominal value to year 2005 real value according to the PCE factors from Bureau of Economic Analysis[33]. Let me call this restricted sample "labor sample". Using "labor sample" I construct the explanatory variable and instrumental variable.

A.1 Efficiency-Adjusted Male Wage Ratio and Technological Change Ratio

According to Mocan and Unel (2011), I construct the efficiency-adjusted labor inputs and earnings. In my labor samples, there are 32 groups (g). Efficiency-adjusted ratio (q) is calculated by the arithmetic mean of the relative wage ratios in the group compared with that in the base group say group 1 (males with less than 12 years of education and less than 10 years of experience) for each group in each metropolitan area in each year. I denote w_{gmt} as the average weekly earnings in each group of each metropolitan area in each year, which is weighted by the product of individual's person weight and weeks worked last year. I denote N_{gmt} as the weighted sum of total weeks worked in that group of that metropolitan area in year t. Then I can get the total efficient adjusted labor supply of high (H_{mt}) and low (L_{mt}) educated workers in each metropolitan area in each year, and also the corresponding wage for high (W_{hmt}) and low (W_{lmt}) educated workers in each metropolitan area in each year using following equations.

(A.1) $H_{mt} = \sum_{g\in G_H} q_g N_{gmt}\ ,\quad L_{mt} = \sum_{g\in G_L} q_g N_{gmt}$

(A.2) $\mathrm{W_{hmt}} = \sum_{\mathrm{g\in G_H}} \frac{\mathrm{w_{gmt}N_{gmt}}}{\mathrm{H_{mt}}}\ ,\quad \mathrm{W_{lmt}} = \sum_{\mathrm{g\in G_L}} \frac{\mathrm{w_{gmt}N_{gmt}}}{\mathrm{L_{mt}}}$

Here, h and l indicates education category, m indicates metropolitan area, and t indicates year. Then, I calculate technological shock ratio denoted as AR=ln(Ah/Al) from equation (2.8) as follows.

(A.3) $\frac{W_{hmt}}{W_{lmt}} = \left(\frac{A_{hmt}}{A_{lmt}}\right)^{\frac{\sigma-1}{\sigma}} \left(\frac{H_{mt}}{L_{mt}}\right)^{-\frac{1}{\sigma}}$

[32] The top-coding values are referred to IPUMS data description and David Autor's code.

[33] "Price Indexes for Personal Consumption Expenditures by Major Type of Product" BEA Table 2.3.4 http://www.bea.gov/national/nipaweb/TableView.asp?SelectedTable=64&ViewSeries=NO&Java=no&Request3Place=N&3Place=N&FromView=YES&Freq=Year&FirstYear=1969&LastYear=2011&3Place=N&Update=Update&JavaBox=no#Mid The relevant value for the PCE is 69.025 in 1989, 87.636 in 1999 and 105.499 in 2007. Use 2007 value PCE for 2007 ACS is consistent with David Autor's code in which he uses 2008 PCE value for 2008 ACS data.

Assume $\sigma = 1.6$. We can get the following:

(A.4) $AR = ln\left(\frac{Ah}{Al}\right) = ln\left(\frac{A_{hmt}}{A_{lmt}}\right) = 2.667 * ln\left(\frac{W_{hmt}}{W_{lmt}}\right) + 1.667 * ln\left(\frac{Hmt}{Lmt}\right)$

I use the same (q) value and apply the similar procedures in (2.5) in calculating male wage gap. I restricted samples to include only men and calculate wage gap between high educated and low educated male workers aged 21 to 40. And finally I use the logarithm of the wage ratio as the key explanatory variable denoted by $\ln(\omega_{MH}/\omega_{ML})_{mt}$.

A.2 Composition-Adjusted Male Wage Ratio and Technological Change Ratio

An alternative measure of earnings is the composition-adjusted male wage ratio. As in Autor et al. (2008), this measure adjusts the male wage ratio according to the proportion of each group's labor input over that of all groups. Hence, the average earnings are not fluctuated by the change in each category, defined by gender, education and experience. Follow Mocan and Unel (2011), the composition index for each group (n_{gm}) is calculated by the arithmetic mean of n_{gmt} over three waves of the surveys in each metropolitan area, where $n_{gmt} = N_{gmt}/\sum_g N_{gmt}$. Then, the composition-adjusted weekly wage is calculated as follows.

(A.5) $W_{hmt} = \frac{\sum_{g \in G_H} n_{gm} w_{gmt}}{\sum_{g \in G_H} n_{gm}}$, $W_{hmt} = \frac{\sum_{g \in G_L} n_{gm} w_{gmt}}{\sum_{g \in G_L} n_{gm}}$.

The rest of the calculation is as same as the efficiency-adjusted ratio.

A.3 Race-Specific Male Wage Ratio

To calculate the race-specific male wage ratio, I divided the sample in each year into 64 groups. Odd groups are for white and even groups are for non-white observations. Thus, the mean wage ratio within each group is defined by the mean wage over all individuals who have the same race, gender, education and experience indicators. And the total weeks worked is also calculated within the newly defined groups.

The technological shock ratio is still defined as the demand ratio among all high and low educated workers. The male wage ratio for all is calculated by all male workers using similar procedures except for the value of within group variables are different, since there are 64 groups instead of 32 groups. However, the white male wage ratio is restricted to only high and low educated white males. The non-white male wage ratio is for only high and low educated non-white males.

Table (A.2) presents the number of observations in each metropolitan area to calculate male wage gap and technological change ratio for cases with 32 groups. It also shows the number of observation in each group within the largest metropolitan area and one of the smallest metropolitan areas included in the sample. I use individuals within each group of metropolitan area to calculate the mean wage ratio of each group, w_{gmt} used in equation (A.5). Typically, sample size is around 370 in each group of a moderate size metropolitan area.

After the variable, I select female aged 21 to 35 in "raw sample" of three years. And I further delete all individuals migrated within last 5 years only in the 1990 and 2000 Censuses, given the information missing in the 2007 ACS. I then merge the three samples respectively with files containing constructed variables.

Finally, I merge samples from these three waves of the surveys into one big panel by metropolitan area index. I construct two race dummies (black/white or black/non-white), three year dummies, 212 metropolitan specific year trends and clusters defined by metropolitan area, year and age, which means all females with same age in the same metropolitan area in the same year are in one cluster.

Table A.1 Observation Deleted Due to Partial Identification in the Metropolitan Areas

Number	Name	Observation deleted		
		1990	2000	2007
0	N/A	4,439,982	4,172,071	823,400
1	Lawrence-Haverhill, MA/NH	6,047	13,012	2,663
2	Bridgeport, CT	11,915	17,304	3,318
3	Brockton, MA	5,520	12,872	2,566
4	Clarksville-Hopkinsville, TN/KY	4,721	5,957	1,431
5	Columbus, GA/AL	0	9,065	1,856
6	Flint, MI	15,994	9,243	1,663
7	Hartford-Bristol-Middleton-New Britian, CT	23,372	35,919	7,155
8	Hickory-Morgantown, NC	4,985	17,691	3,784
9	Houma-Thibodouw, LA	8,495	4,245	933
10	Jackson, MS	10,617	18,813	4,601
11	Johnson City-Kingsport-Bristol, TN/VA	13,299	14,914	3,220
12	Lafayette, LA	7,546	11,359	2,245
13	Lexington-Fayette, KY	8,318	10,066	2,470
14	Macon-Warner Robins, GA	7,149	16,778	3,376
15	Manchester, NH	4,301	3,874	824
16	Nashua, NH	3,820	4,044	878
17	New Bedford, MA	6,599	9,148	1,840
18	New Haven-Meriden, CT	9,429	16,446	3,371
19	New London-Norwich, CT/RI	6,253	0	0
20	Providence-Fall River-Pawtuckett, MA	12,043	50,107	9,943
21	Pawtucket-Woonsocket-Attleboro, RI-MA	5,517	0	0
22	Reno, NV	5,937	17,297	4,005
23	Roanoke, VA	5,119	11,423	2,431
24	St. Cloud, MN	7,178	10,017	2,116
25	Salt Lake City-Ogden, UT	30,163	54,438	13,767
26	Springfield, IL	10,230	4,739	1,070
27	Springfield-Holyoke-Chicopee,MA	17,759	28,508	5,766
28	Stamford, CT	4,579	16,797	3,478
29	Waterbury, CT	4,657	5,182	1,019
30	Worcester, MA	14,192	13,917	2,939
	Total	4,715,736	4,615,246	918,128

Note: Code "N/A" represents the unidentified individuals in the metropolitan area which is more than 70% identified.

Table A.2 Observations to Construct Instrument and Main Explanatory Variable in 1990 Sample (Selected)

Metropolitan area	Freq.	Percent	Cum.
Abilene, TX	1,794	0.07	0.07
Akron, OH	8,705	0.35	0.42
Albany-Schenectady-Troy, NY	13,539	0.55	0.97
………			
Yuma, AZ	1,308	0.05	100.00
Total	2,480,945	100.00	

Within MA

Groups defined by sex, education and experience	New York-Northeastern NJ (Biggest MA, 9.08%)			Yuma, AZ (Smallest MA, 0.05)		
	Freq.	Percent	Cum.	Freq.	Percent	Cum.
1	1,837	0.82	0.82	27	2.06	2.06
2	3,057	1.36	2.17	54	4.13	6.19
3	3,180	1.41	3.58	31	2.37	8.56
……						
32	3,361	1.49	100	10	0.76	100
Total	225,388	100		1,308	100	

Vita

Xue Li was born in Panshi City, Jilin Province in China. She earned her Bachelor of Science in Economics from Peking University in China in 2008 and her Master of Science in Economics from Louisiana State University in the U.S in 2010. She pursued her doctorate degree in Economics under the supervision of Dr. Naci Mocan at Louisiana State University. She also worked as a teaching assistant in the Department of Economics and received full assistantship during her graduate studies. She will join the Center for Economics, Finance and Management Studies at Hunan University in China in the fall of 2014 as a tenure-track Assistant Professor in Economics. Her research interests include Applied Microeconomics, Population Economics, Labor Economics and Health Economics.