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Eric D Gould and Marco Daniele Paserman

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Eric D Gould, Hebrew University of Jerusalem and CEPR
Marco Daniele Paserman, Hebrew University of Jerusalem

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Centre for Economic Policy Research
90–98 Goswell Rd, London EC1V 7RR, UK
Tel: (44 20) 7878 2900, Fax: (44 20) 7878 2999
Email: cepr@cepr.org, Website: www.cepr.org

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ABSTRACT

Waiting for Mr Right: Rising Inequality and Declining Marriage Rates*

This Paper demonstrates that women search longer for their first or second husband in cities with higher male wage inequality, and analyses several explanations for this result. A causal link is established by showing that the results are robust to the inclusion of city fixed-effects and city-specific time trends, and by using inequality in the woman's state of birth as a proxy for the local level of male inequality. Increasing male inequality explains about 30% of the marriage rate decline for women over the last few decades, and we show that this is not due to the effects of male inequality on female inequality, female labour supply, or decisions by women to re-locate, nor to the decisions of men in reaction to changes in their own wages. The evidence supports the idea that higher male inequality increases the option value for single women to search longer for a husband.

JEL Classification: D31, J12 and J31
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Eric D Gould
Department of Economics
Hebrew University of Jerusalem
Mount Scopus
91905 Jerusalem
ISRAEL
Tel: (972 2) 588 3247
Fax: (972 2) 581 6071
Email: mseric@mscc.huji.ac.il

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Marco Daniele Paserman
Department of Economics
Hebrew University of Jerusalem
Mount Scopus
91905 Jerusalem
ISRAEL
Tel: (972 2) 588 3365
Fax: (972 2) 581 6071
Email: dpaserma@shum.huji.ac.il

For further Discussion Papers by this author see:
www.cepr.org/pubs/new-dps/dplist.asp?authorid=152194

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1. Introduction

This paper examines the connection between male wage inequality and female marriage rates within cities. This relationship might exist for several reasons. An increase in male inequality could be considered an increase in the dispersion of husband quality, and thus may increase the option value for women to search longer for a husband. However, higher inequality could also cause women to invest more in their careers in order to exploit higher returns to skill, thus causing delays in marriage. Also, higher male inequality may be correlated with increasing female inequality, which may, in turn, cause men to search longer for a wife. In addition, higher male inequality may be correlated with lower marriage rates due to a tendency for single women to move to cities with higher inequality. Finally, changes in male inequality may affect marriage rates simply due to the decisions of men in response to changes in their own wages. For example, as rich men get richer they may search less (due to higher search costs) or become more selective, and as poor men get poorer they may be less attractive in the marriage market. Thus, inequality may be causing declines in marriage simply by increasing the proportion of men at the tails of the wage distribution where marriage typically occurs at an older age.

The goal of this paper is to establish whether a causal connection exists between inequality and marriage rates, and then to examine the evidence for the various explanations suggested above. Our empirical strategy is to explain the individual marital decisions of white women using data from the 1970, 1980, and 1990 Censuses. The use of individual level data allows us to control for a variety of personal characteristics (age, education, etc.) and city-level variables that characterize the local labor market and marriage market (the mean male wage, mean female wage, sex ratio, population, etc.). After controlling for these variables, we identify the effect of male wage inequality on women's marriage decisions by exploiting city-level variation in the distribution of male wages that each woman faces. We find a strong positive relationship between male wage inequality and the probability of a woman being single. This basic finding applies to women searching for their first or second husband, and is not limited to a particular age or education group. Furthermore, the results are shown to be robust to the use of various measures of male wage inequality, and to the inclusion of city fixed-effects and city-specific time trends, implying that *changes* in inequality are correlated with *changes* in marriage rates. Our causal interpretation is further supported by using the exogenously determined level of inequality in the woman's state-of-birth as a proxy for her local level of male inequality.

The results suggest that increased inequality may account for up to 30% of the overall decline in female marriage rates in the last few decades. We show that this result is not due to women working harder in cities with higher male inequality, and is not due to single women moving to cities with higher male inequality. We also demonstrate that women delay their marriage decisions only in response to higher local male inequality, not higher female inequality. Furthermore, we show that men also delay marriage when there is higher male inequality, but that they do so after conditioning upon their wage and rank in their city's wage distribution. That is, the same man with the same wage and rank gets married later in a city with a higher level of male inequality. Thus, higher inequality is clearly not causing delays in marriage due to the responses of men to changes in their own wages. Overall, the evidence is supportive of the idea that higher male inequality increases the option value for women to search longer for a husband. This is the only explanation consistent with the robustness of these results to women of all age and education groups and to women searching for their first or second husband.

This paper is related to the recent literature on the consequences of higher local levels of inequality. For example, higher inequality has been linked with lower investments in social capital, falling labor force participation rates for men, higher crime rates, higher local price levels, and higher investments in education.[1] Although there are several reasons why male inequality may affect a city's marriage patterns, the recent literature has so far concentrated on the reverse -- how assortative mating patterns are affecting aggregate levels of family income inequality (Kremer, 1997; Fernandez and Rogerson, 2001; Greenwood, Nezih, and Knowles, 2000). The research most closely related to ours has focused on explaining the age at first marriage. Goldin and Katz (forthcoming) provide evidence that oral contraception has caused women to delay marriage. Bergstrom and Schoeni (1996) argue that the income of males is positively associated with their age at first marriage, while Blau, Kahn, and Waldfogel (2000) show that women delay marriage when their own labor market prospects improve or when the male labor market deteriorates. This paper incorporates many of these same factors, but differs by focusing on the various reasons underlying a causal relationship between inequality and marriage (and re-marriage).

The rest of the paper is organized as follows. The next section outlines the various explanations for the relationship between inequality and marriage. Section 3 presents the data and the econometric methodology. The main results are analyzed in Section 4 and Section 5 addresses alternative explanations for our results. Section 6 concludes the discussion.

[1] See Alesina and La Ferrara (2000); Juhn (1992); Gould, Weinberg, and Mustard (2002); Frankel and Gould (2001); and Gould, Moav, and Weinberg (2001) respectively.

2. The Link Between Inequality and Marriage

This section presents several explanations for the link between inequality and marriage, and discusses the empirical implications of each one.

Women Searching for Men. A simple marriage search model consists of women receiving marriage offers from men in each period with some probability. If a woman accepts the offer, she will enjoy the utility associated with it in future periods. If she rejects the offer, she enjoys the current value of being single and the option value of potential marriage offers in the future.[2] In this set-up, which is analogous to the prototypical one-sided search model of Lippman and McCall (1976), it is well known that a mean-preserving spread in the quality distribution of offers leads to increased selectivity in the offer acceptance strategy (Burdett and Ondrich, 1985). The intuition for this result is straightforward. In the extreme case of a degenerate offer distribution, the solution to the search problem is trivial: if the value of being married exceeds the value of being single, the woman will accept the first marriage proposal she receives, as there is no value to waiting. As the offer distribution spreads out, it is no longer optimal to accept any marriage proposal: declining a low quality proposal in the current period opens up the possibility to draw a potentially better offer in the next period. In this case, waiting has an option value, which increases with the variance of the offer distribution. Therefore, higher dispersion of husband quality is expected to lead to delays in the marriage commitments of women, due to the increasing option value of waiting for potentially higher offers. Hence, if the wage is a non-negligible

[2] The idea of using this search theory approach dates back to Becker (1973, 1974), and has also been successfully used to analyze marital dissolutions (Becker, Landes and Michael, 1977; Weiss and Willis, 1997).

attribute of a potential husband's quality, an increase in male wage inequality should be associated with a decrease in marriage rates.[3]

Men Searching for Women. Clearly, the same search model can be applied in reverse: if women make marriage offers and men can decide whether to accept or reject them, we should expect marriage rates to be negatively correlated with female wage inequality. However, given the traditional division of labor within the household, variation in market wages may play a smaller role in determining variation in the quality of women. Therefore, we suspect that the effect of female wage inequality on the decisions of men is smaller than the effect of male wage inequality on the decisions of women.

Women Exploiting Higher Returns to Education, Experience and Skills. A link between inequality and marriage rates may exist independently of any strategic games in the marriage market. In response to higher inequality, women may increasingly invest in education and work experience in order to exploit the higher returns to these activities, and therefore, delay marriage. Alternatively, risk averse women may work more to hedge against the increased riskiness of their partners' wages. Two testable implications of this model are that female inequality should be negatively correlated with marriage rates, and positively correlated with female labor supply.

Signaling Model of High Ability Men. Bergstrom and Bagnoli (1993) propose a model in which high ability men delay proposing marriage until their wages accurately signal their ability level. They argue that waiting allows some of the uncertainty surrounding the quality of high ability men to dissolve. If high wage

[3] It is conceivable that, holding constant other attributes (sense of humor, caring, physical attractiveness, shared interests, etc.) a high wage partner will be preferable to a low wage partner because of the higher consumption level enabled by the higher joint family budget constraint. This result relies only on the assumption that consumption is a normal good and that the distribution of family income (the bargaining power of each partner) is not affected by income.

inequality reflects higher uncertainty about a man's true quality, then we should expect a negative relationship between inequality and marriage rates. However, the predictions of this model are not unambiguous, since higher inequality exacerbates the differences between men, which perhaps makes the signal even clearer at an earlier stage. The model does, however, produce the empirically testable prediction that high wage men marry later than low wage men.

Men Reacting to Changes in their Own Wages. If the marriage decisions of men are influenced by their own wages, then increasing inequality can affect the overall marriage rate simply by changing the proportion of men with any given wage level. For example, rich men may be more selective in their choice of partners, and therefore, wait longer to get married. Poor men may also take longer to get married since they are deemed less "marriageable." Under this scenario, men with wages closer to the mean are more likely to get married earlier. Therefore, a widening distribution of male wages could lead to lower marriage rates simply because rich men are getting richer and poor men are getting poorer. That is, men are just reacting to their own wages, and consequently, the level of inequality is positively associated with the overall local marriage rate. However, after controlling for a man's own wage, we expect that the local level of male inequality should no longer affect marriage decisions.

In the empirical work that follows, we assess the evidence for each of these theories by examining their testable implications.

3. The Data and Empirical Methodology

To establish the link between male wage inequality and marriage patterns, we exploit variation in inequality and marriage rates across metropolitan areas and over time. The basic assumption underlying our identification strategy is that the metropolitan area can be taken as the local marriage market of reference.[4] The data is taken from the 1970, 1980, and 1990 1% Public Use Micro Samples (PUMS) of the United States Census. The large samples enable us to construct accurate measures of inequality in each metropolitan area (MA) and in each sample year.

Our basic measure of inequality is the standard deviation of log weekly wages for full-time male employees, aged 16 to 64, who were not in school and who worked at least one week in the calendar year prior to the census year. This measure captures the price of skill, and therefore, can be considered an accurate proxy for the permanent component of the male earnings distribution. In the empirical analysis, we match each woman to the measure of overall inequality for the entire male population in her metropolitan area, therefore disregarding assortative mating on the basis of age and education. Two reasons motivate this choice: first, changes in inequality may affect women's schooling decisions, thus casting doubt on the exogeneity of inequality in a given education group; second, data constraints make it difficult to construct accurate local inequality measures within finely defined cells.

The sample in our main analysis consists of over 300,000 white women between the ages of 21 and 30 in 321 metropolitan areas in the three census years.[5] Sample

[4] This seems reasonable based on the fact that the overwhelming majority of the population does not move frequently across metropolitan areas: of those aged 20 to 30 – the most mobile age group – who lived in a metropolitan area in 1975 and 1985, 73% and 75% were living in the same metropolitan area in 1980 and 1990 respectively. These percentages are obviously much higher for other age groups. Source: Authors' calculations based on the PUMS.

[5] The restriction on white women only is driven by both theoretical and data considerations: on one hand, because of the low rate of interracial marriages it is appropriate to analyze separately the marriage markets for African Americans and for whites; on the other hand, it is difficult to estimate precise measures of inequality among African American men at the metropolitan area level using the 1% PUMS due to very small sample sizes. In the smallest metropolitan area, (Columbia, MO) the male inequality measure was based on a sample of 140 observations.

statistics for each year are shown in Table 1 and Figures 1 and 2, which exhibit the basic trends of increased wage inequality and declining marriage rates.[6]

Figure 3 displays the strong correlation between inequality and marriage rates across metropolitan areas within each sample year. The simple regression coefficient is statistically significant and increases over time: the slope coefficient is 0.217 in 1970 and 0.530 in 1990. This simple correlation could be due to unobservable factors across cities that are correlated with both male inequality and the proportion of unmarried women. Figure 4 controls for these unobservables by regressing the changes in the proportion of single women on changes in male inequality within the city over time. Across each pair of years, inequality is still a very significant determinant of the change in the marriage rate, thus implying that heterogeneity in fixed unobservable city characteristics is not responsible for the correlations seen in Figure 3.

This first pass through the data is highly suggestive that higher inequality causes delays in marriage. However, the results in Figure 4 could be due to a correlation between city level changes in inequality and unmeasured city level changes in labor market and marriage market conditions. Therefore, the goal of the rest of the paper is to control for this possibility, and establish a causal connection between higher inequality and delayed marriage. To do this, we adopt an empirical strategy that controls for other changes in local labor market and marriage market conditions, and exploits the micro-level data of the Census to model the individual marital choices of women.

[6] There is an extensive literature documenting the increase in male wage inequality over the last four decades: see Bound and Johnson (1992), Katz and Murphy (1992), Juhn, Murphy, Pierce (1993), Gottschalk and Moffitt (1994), and Gould (2002).

The basic empirical specification for our regressions is a probit model of the woman's decision to remain single:

$$P(y_{ijt} = 1) = \Phi(\alpha' Ineq_{jt} + \beta' Z_{ijt} + \gamma' X_{ijt} + \eta_{jt}), \qquad (1)$$

where y_{ijt} is a dummy variable equal to 1 if woman i living in metropolitan area j at time t has never been married and 0 otherwise; $Ineq_{jt}$ is a measure of inequality among white men in metropolitan area j at time t; Z_{ijt} is a vector of controls for marriage market and labor market conditions faced by woman i; X_{ijt} is a vector of individual characteristics; η_{jt} is a city-year specific effect capturing unobservable aspects of the local marriage market, whose exact specification will be discussed later. Our main focus of interest is the parameter on city-level inequality α: according to the search theory model described in Section 2, we expect α to be positive.

Marriage market conditions are proxied by total city population and by the sex ratio, defined as the ratio of total white men over total white women in the metropolitan area.[7] The former captures the probability that a woman meets a potential partner, while the latter measures the extent to which women in a given city compete for the same restricted pool of men.[8] To proxy for labor market conditions, we include the mean hourly wages of men and women in the metropolitan area, and an index of relative local demand for female labor.[9] The relative demand index follows Katz and Murphy (1992) and Gould, Weinberg, and Mustard (2002), and is based upon the initial local industrial composition and national industrial shifts which may favor female workers relative to male workers (see Blau, Kahn and Waldfogel,

[7] These are calculated using the personal weights provided by the Census.
[8] See Angrist (2002) for an extensive study on the importance of sex ratios in explaining marriage patterns.
[9] Hourly wages are used for two reasons. First, hourly wages are a better measure for women since women often work part-time. Second, we use hourly wages for men so that when we include both measures for men and women, the difference can be interpreted as the gender gap in hourly wages.

2000, for a similar strategy). The regressions also include three educational attainment dummies and a full set of age dummies. Sample statistics for selected variables are presented in Table 1.

4. The Effect of Male Inequality on Female Marital Status

Table 3 presents the estimates for the basic specification of our model under three alternative assumptions about the structure of the unobserved city-year effect. We first estimate the model including only aggregate year dummies, omitting any city-specific fixed-effects (formally, we let $\eta_{jt} = \delta_t$). The year dummies capture all the national trends in the tastes for marriage: social norms, technological progress in household production, changes in divorce laws, contraceptive methods, etc. In this specification, we essentially pool together all three years of data, and exploit both cross-sectional and time series variation in inequality to estimate the parameter of interest.

To account for possible unobserved city characteristics that are correlated with the propensity to stay single, we also estimate the model with city fixed-effects (formally, $\eta_{jt} = \theta_j + \delta_t$). This strategy identifies the parameter of interest from variation in the city-level time series in inequality and marriage rates, and yields consistent estimates if the unobserved metropolitan area characteristics correlated with marriage and inequality are indeed constant over time. The inclusion of city fixed-effects could be considered an overly conservative identification strategy because we are throwing away all of the strong cross-sectional variation in marriage rates and inequality, as seen in Figure 1.[10]

[10] See Hamermesh (2000) for a discussion of this issue.

Even with the inclusion of city fixed-effects, we could still obtain a spurious correlation if *changes* in omitted city-level variables are correlated with changes in inequality and marriage rates. We cannot resolve this problem by including a dummy variable for every possible city-year combination in the data, since this would saturate the regression (i.e., there would be zero degrees of freedom in the estimation). Instead, the city-year effect is modeled as the sum of a city fixed-effect, an aggregate time effect, and a city-specific linear time trend: $\eta_{jt} = \theta_j + \delta_t + \lambda_j t$. Again, this is a conservative estimation strategy, since it throws away all cross-city variation in the levels and trends in inequality. The parameter of interest is identified only from deviations in male inequality and in marriage rates from a city-specific linear trend (i.e. whether the within-city acceleration of male inequality explains the within-city deceleration of marriage rates, after controlling for the other control variables). Note that this estimation strategy is feasible only if metropolitan areas are observed in all three sample years.[11] Therefore, the sample in this third model is restricted to only those metropolitan areas that were consistently defined in all three Census years.

Table 3 presents the results for the three main specifications of the model, using the standard deviation of log male wages as the measure of male inequality. In all specifications, the standard errors are corrected to take account of both the grouped structure of the error term, and the fact that the main variable of interest (inequality) varies at a higher level of aggregation than the individual units. Without this correction, conventional standard errors may be severely underestimated (Moulton, 1986).

[11] This strategy also uses up most of the effective degrees of freedom in the regression. Although we are using individual-level data, the measures for city-level inequality vary only by city and for 3 years. Including a dummy for each city and each city's time trend, therefore, means that the number of coefficients is at least two-thirds of the number of observations for city-level inequality.

The table shows that after controlling for the education of the woman and her local labor market and marriage market prospects, the probability that a woman is single increases with higher local male wage inequality. This result is in accordance with the theories described in Section 2. The coefficient on male inequality is statistically significant in all three specifications. Interestingly, the magnitude of the coefficient increases as we include city-specific fixed-effects and a city-specific time trend. This suggests that, maybe contrary to expectations, the unobserved preference for staying single is negatively correlated with male inequality.

In general, the control variables are statistically significant and have the expected sign: marriage rates decline with higher education, higher wages for women, and demand shifts in favor of women; marriage rates increase with age, higher wages for men and a higher ratio of men to women. Overall, the results show that women get married less when their labor market prospects improve (relative to men), and they get married more when marriage market conditions improve and when labor market prospects for men are relatively better.[12]

Not only is local male inequality a statistically significant determinant of the marital status of each women, the size of the effect is economically meaningful. Table 3 presents how much of the marriage rate decline can be predicted by the increasing inequality trends using the estimated marginal effects in Table 2. For the change between 1970 and 1990, inequality can explain between 18.4 percent and 28.8 percent of the decline in the marriage rate of 21-30 year old women. The larger estimate comes from the specification with the city-specific time trends, lending support to the hypothesis that the link between male inequality and marriage is indeed causal. The estimates for the other time periods are similar, although inequality seems

[12] These findings are similar to those of Blau, Kahn, and Waldfogel (2000).

to explain more of the 1980-1990 decline in the incidence of marriage than the decline between 1970 and 1980.

Table 4 checks the robustness of these results using alternative measures for male inequality and samples. Having ascertained that unobserved changes in the propensity to marry are unlikely to bias the results upwards, we restrict attention to the pooled specification and to the specification with only metropolitan area fixed-effects. We first check whether the results in Table 3 are due to differences in the age and education composition of the male workforce across metropolitan areas, by using two alternative measures of residual inequality. Both measures control for inequality *between* age and education groups. The second measure also controls for the marital status of men in the MA, since it could be argued that lower marriage rates cause higher inequality because married men earn more than unmarried men. Therefore, the second measure of residual wages eliminates this difference. The results for both measures of residual male inequality are very similar to those for overall male inequality, which is not surprising considering that most inequality within men is due to inequality *within* groups and not *between* groups (see Juhn, Murphy, and Pierce, 1993).

Previous empirical work has shown that single men tend to engage in risky behavior (Akerlof, 1998). If the erratic behavior of single men translates into higher inequality *within* single men, it could produce a spurious correlation between declining marriage rates and overall male inequality. To eliminate the effect of inequality *within* single men, the next measure of inequality is constructed using only a sample of married men. The coefficient estimate is still positive and significant, thus showing that our basic findings are not due to the effect of unmarried men on the local level of inequality.[13]

Table 4 next shows that alternative measures of inequality, such as the Gini coefficient and the 90-10 log wage differential, also yield similar results to Table 3. Looking only at inequality in the top half of the distribution, the 90-50 log wage differential has a strong effect on marriage rates in the pooled model, but the effect goes away when city effects are included. The opposite is true for the 50-10 log wage differential.

Table 4 also presents the probit results for regressions run separately for different age and education groups of white women. For the youngest age group (ages 16-20), the results are not significant, which is not surprising since factors other than inequality are more likely to be driving the marriage decisions of very young women (investments in education, lack of contact with wage-earning men, unexpected pregnancies, etc.). In contrast, results for the older group of women (ages 31-35) are statistically significant. Although the coefficient is smaller in magnitude than that for the middle age group (ages 21-30), the change in marriage rates for the older age group is also smaller, since a high proportion of women is already married by the age of 30. Consequently, the percentage change in marriage rates for the older group between 1970 and 1990 attributable to changes in inequality is very similar to that found for the 21-30 age group (33.6 percent versus 25.4 percent for the middle age group).[14]

[13] If there is positive sorting on ability into marriage, then inequality within married men should be positively associated with increasing marriage rates. Thus, the estimate on inequality for married men is likely to be biased downwards: higher inequality should be negatively related to the propensity to be single.

[14] The number for the older group is calculated by multiplying the marginal effect found in Table 4 (0.128) by the change in inequality between 1990 and 1970 (0.64 - 0.49), and dividing by the change in marriage rate (0.94 – 0.88) for the older group.

The bottom two rows of Table 4 show that the results are also very similar for educated and less-educated women. Since there is a high degree of assortative mating in the marriage market, this last result proves that the effect of male inequality on marital status is clearly not exclusive to less-educated women who may be finding it harder to find a suitable less-educated husband (whose wages are declining), nor is it exclusive to educated women who are trying to find an educated husband (whose wages are increasing).

Table 5 examines whether women who go through a divorce also take longer to get re-married in response to higher inequality. This may already allow us to discriminate between alternative theories: for example, search theory does not discriminate between women who have never married and women who are single again after divorce: the option value of waiting to select a partner should be higher for both types of women in the presence of higher local male inequality.

For samples of older and younger groups of women who have experienced a divorce, Table 5 presents our basic probit specification for whether a divorced woman is still single or not.[15] The results indicate that divorced women indeed wait longer to get re-married when there is higher male inequality. The marginal effect of higher inequality is roughly equivalent to what was found with women looking for their first husband in the previous tables (about 0.30). In addition, Table 5 presents the results for a "Heckman selection-corrected" model (Heckman, 1979), which attempts to correct for the non-random selection of women into the divorced sample. The Heckman model includes a first stage probit on whether the woman experienced a divorce, and a second stage probit for whether the divorced woman is still single. The

[15] Due to the fact that the 1990 Census does not identify whether a married woman is on her first or second marriage (as do the 1970 and 1980 Censuses), data from 1990 are dropped for our analysis on re-marriage.

first stage probit does not contain variables excluded from the second stage, so identification may be weak. With this caveat in mind, the selection-corrected model is used to analyze the sensitivity of the simple probit model. The results for both models presented in Table 5 are almost identical, so sample selection appears not to be an important factor. Altogether, these results show that the effect of male inequality on the decision to delay marriage is not limited to women who are searching for their first husband.

Overall, the results in this section show that higher local male inequality has a statistically and economically meaningful impact on the marital status of a woman. This result is robust to the inclusion of city fixed-effects, city-specific time trends, and various measures of male inequality. Furthermore, this basic finding is not limited to women searching for their first husband, nor is it limited to a certain age group or education group: all mature women are waiting longer to get married when faced with a higher variance of potential husband quality.

5. Interpreting the Results: Potential Explanations

The previous section established the strong relationship between male inequality and female marital status. The findings are strongly supportive of a search model of the marriage market. Higher inequality raises the option value of being single, and therefore causes delays in marriage. This theory should apply equally to women of all age and education groups, and to women in search of their first or second husband, and this is indeed what we find. Moreover, the effect of the other control variables (see Table 2) on the propensity to stay single is also consistent with a search model: women are more selective when the value of being single increases (higher female wages), less selective when the value of being married increases (higher male wages), and more likely to find a husband when the probability of receiving an offer increases (a higher sex ratio).

The rest of this section examines the alternative theoretical explanations (described in Section 2) for the relationship between inequality and age at first marriage, and assesses the empirical evidence for each one.

The Effect of Female Inequality

If there is a strong correlation between female inequality and male inequality, the results in the previous section could be picking up the effect of female inequality on the marital status of each woman. This could be due to men waiting longer to choose a wife because the quality distribution of potential wives is spreading out, or because the competitive environment in the marriage market changes in such a way so that it takes women longer to settle down.

Analyzing this issue has one important empirical problem – measuring female wage inequality is handicapped by the lack of information on a woman's actual work experience. This problem is much bigger for women than men since women tend to work less than men. We tackle this problem by working with residual female wage inequality. Using residual female inequality will at least control for the selection of female workers in the labor force based upon their observable characteristics (age and education).

Table 6 shows that female residual inequality actually has a negative impact on the probability of being single. This result is true even after controlling for male residual inequality. But, more importantly, the effect of male residual inequality is still positive and significant. The inclusion of female residual inequality does not affect this result. Regarding the negative coefficient on female residual inequality, we suspect that this is due to reverse causation: in areas where women tend to marry more, women will also tend to work less, thus raising our measure of female inequality due to the higher variance in unobserved work experience. Another possible explanation is that higher male inequality raises the bargaining power of women inside the household (since they are more inclined to reject offers and keep searching for a better offer), thus decreasing their labor supply (see Chiappori et al., 2002; Grossbard-Schechtman, 1993; Angrist, 2002).

In Table 7, we directly address the issue of whether women work more in areas with higher male inequality, conditional on their marital status and number of children. As mentioned in Section 2, high work effort may yield larger returns in cities with higher inequality. Table 7 examines this issue with two different model specifications. The first model is a probit for whether the woman works full-time (at least 30 hours a week), and the second model is a generalized Tobit model on weeks worked in the calendar year prior to the census.[16] Each specification includes MA fixed-effects and the other control variables used in previous analyses.

The results in Table 7 reject the hypothesis that women work more when there is higher male inequality. This result is robust to both models and to the inclusion of female residual inequality. Therefore, there is no evidence that our main results are due to women working more when there is higher local male inequality. Furthermore, Table 7 supports our interpretation that the coefficient in Table 6 for female inequality is due to reverse causation: Table 7 clearly shows that higher female inequality is associated with women working less – thus causing us to attribute variation in actual female experience to our measure of female wage inequality. Consequently, when

[16] The generalized Tobit model accounts for the fact that in 1970 data on weeks worked comes only as a grouped variable.

women get married more, they work less and cause our measure of female wage inequality to increase (see Table 6).

Do Single Women Move to Areas With Higher Male Inequality?

If single women tend to move to areas with higher wage inequality, this endogenous moving could be responsible for the positive relationship between male inequality and the tendency to remain single. However, it should be noted that even if this were the case, it would still support our search interpretation – it would just mean that women are employing costlier search methods (moving between cities) than we anticipated. If, however, single women just happened to be moving to cities with higher inequality for reasons unrelated to searching for a husband, then our results would be spurious. We address this issue by controlling for whether a woman has moved between states in Table 8.

The first column in Table 8 repeats our basic probit analysis of female marital status using male inequality in the woman's current state of residence, rather than the woman's current MA of residence. The first column also controls for state of residence fixed-effects and shows that our basic result is robust to using state level measures for local male inequality. The second column in Table 8 controls for whether the woman has moved states since her birth (i.e. her current state of residence differs from her state of birth) and shows that the results are unaffected by controlling for endogenous moves between states. The third column controls for state fixed-effects both for the current residence and for the residence at birth. That is, we control for the state where she was born and where she lives at the time of the survey. Again, the results are unchanged. Finally, the last column shows that the results are robust to the use of male inequality in the respondent's state of birth instead of her current state. The use of inequality in the state of birth provides strong evidence of the causal interpretation of our basic results: it is impossible for a woman's choice of marital status to be endogenous to the level of male inequality in her state of birth, since there is no possible choice that she can make that can affect where she was born. Overall, Table 8 shows that the main results are not due to single women moving to areas with higher inequality.[17]

Are Men Deciding to Wait Longer When There is More Male Inequality?

In most cases, women marry men, so it should not be surprising that when women delay marriage, men tend to follow suit. However, the connection between higher male inequality and declining female marriage rates could be due to the joint phenomenon of men at the high end of the wage distribution choosing to delay marriage, and of men at the low end becoming "unmarriageable." In this sub-section, we address the issue of whether the decisions of men are driving our results with several different arguments.

First, we show that a probit analysis explaining the marriage status of men yields similar results to our previous analysis on women. The first two columns of Table 9 show that the probability of a man being single is unaffected by the local level of female inequality, but is positively and significantly affected by the local level of male inequality. These results reject the idea that men are waiting longer to get married in response to higher female inequality, and are simply a mirror-image of the results for women: both sexes are waiting longer to get married when there is higher male inequality. In addition, the coefficients on the other variables (not shown) indicate that men get married more when their labor market prospects improve, i.e.,

[17] This is consistent with Costa and Kahn (2000), who find that there were no significant trends between 1970 and 1990 in the locational choices of single women and single men.

when their gains to marriage are *smaller*. [18] These results suggest that marriage decisions are determined by the timing of the woman's preferences rather than the man's preferences. This notion coincides with the common belief that changes in women's preferences for work, education, and marriage have been the main factors driving the changes over time in marriage rates.

More concrete proof is provided in the remaining columns of Table 9. As mentioned above, higher male inequality could be causing delays in marriage by causing rich men who are getting richer to search more and/or become more selective, and by causing poor men who are getting poorer to become "unmarriageable." This hypothesis says that inequality is changing the distribution of men's wages, and the overall marriage rate is declining because the behavior of men with wages at the tails of the distribution is different than the behavior of men towards the center of the distribution. In other words, marriage rates are not declining given the wage of a man, it is just that the wages of men are spreading out and changing the aggregate marriage rate.

In the third column of Table 9, this hypothesis is tested directly by including the residual wage (controlling for age and education, which are already in the regression) of each man as an explanatory variable in the probit. Men without wages are dropped from the regression.[19] The effect of the individual man's wage is negative, which shows that men with higher wages are more likely to get married. This is consistent with the results mentioned above: men who seemingly have less to gain from marriage and more to offer, and therefore, should be more selective in their choice of

[18] These results are not due to endogeneity of marriage and labor market conditions since this results holds true for the relative demand index which is based upon variation (the initial industrial composition and the national trends) which should be exogenous to changes in marital preferences.
[19] Men in school or who did not meet the hours or weeks-worked criteria were dropped from the regression. However, the results with this self-selected sample only confirm the robustness of our basic results to this new sample selection criterion.

spouse, actually are more likely to be married. In addition, this result is inconsistent with the signaling model, which predicts that men with higher wages wait longer to get married so that their signal becomes more clear.

More importantly, the coefficient on male inequality is unaffected (actually a bit larger) after including the man's own wage in the specification. This result is also true after including both the man's residual wage and the residual wage squared – which captures the non-linearity in the hypothesis mentioned above (poor men cannot find anyone to marry and rich men are more selective, while the middle wage men are the most likely to get married). The estimates confirm this non-linearity hypothesis, but the effect of local male inequality is still significantly positive, and the magnitude is unchanged.

The last column in Table 9 includes the man's rank in the distribution of residual wages in his MA. The rank variable controls for the man's standing versus the other local men in competition for the same women, by essentially controlling for how many men are better (or worse) than he is in the local wage distribution. Interestingly, the rank coefficient wipes out the effect of the residual wage but not the residual wage squared, and is itself very significant: the "best" men (the higher ranked men) are clearly more likely to be married, once again illustrating the idea that females are determining the marriage decision. More importantly, the coefficient on local male inequality is unaffected. These results say that a man with the same wage and the same rank in the local wage distribution is more likely to be single if there is more male inequality in his city. Clearly, these results reject the notion that marital status given the wage of the man is unaffected by the local level of male inequality. The local level of male inequality is obviously changing the nature of the local marriage market, and not just the behavior of rich men who are getting richer and poor men who are getting poorer.

6. Conclusion

This paper shows that higher male inequality in a city lowers the marriage rate of women. This basic finding applies to women searching for their first or second husband, and is not limited to a particular age or education group: all mature women are waiting longer to get married when faced with a higher variance of potential husband quality. The result is robust to various econometric specifications and to different measures of local male inequality, including the exogenously determined level of inequality in the woman's state-of-birth. Increasing male inequality explains about 30 percent of the marriage rate decline for women over the last few decades.

Several explanations for these results are examined, but the evidence is inconsistent with many of them. We show that male inequality is not simply acting as a proxy for female inequality, and there is no evidence that inequality lowers the incidence of marriage through an increased work effort channel. Furthermore, the results cannot be explained by the endogenous moving of single women to areas with higher inequality. The results also show that higher wage men get married earlier rather than later as predicted by the signaling model. Finally, the effect of male inequality on marital status is not due to changes in the behavior of men caused by the wages of rich men getting richer and poor men getting poorer. Even conditional upon the wages and rank of each individual man, higher male inequality is clearly altering the fundamentals of the local marriage market, resulting in lower marriage rates. Overall, the evidence is supportive of the idea that higher male inequality increases the option value for women to search longer for a husband. This is the only explanation consistent with the robustness of these results to women of all age and education groups and to women searching for their first or second husband.

Overall, this paper makes at least three important contributions. First, the results increase our understanding of the factors that determine a city's marriage rate. Second, the analysis determines the degree to which inequality has contributed to the overall decline in marriage rates in the United States in the last few decades. Third, we identify a new consequence of the growing disparity in the wage distribution – declining marriage rates. Although the causes of higher inequality have been studied extensively, the consequences on society have received less attention. Here, we demonstrate the effect of inequality on declining marriage rates, which may have important implications for society. For example, higher rates of single men and women may lead to more out-of-wedlock children, loss of the utility from marriage, excessive search in the marriage market, higher crime, lower incentives to work hard, etc. A possible extension of this analysis would examine the role of inequality on divorce rates, but the search theory predictions regarding divorce are not unambiguous: inequality raises the incentive to search outside the marriage, but since women tend to marry high wage men, their gains from the current marriage may also be higher. Sorting out all these issues will allow for a more complete assessment of the impact of higher inequality on society as a whole.

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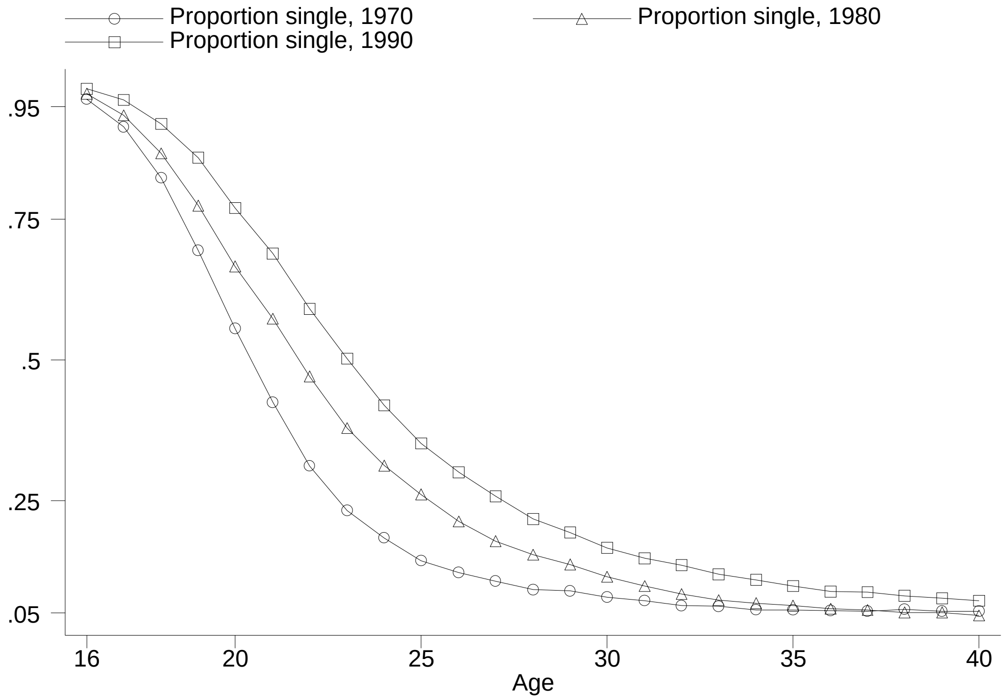

Figure 1: Proportion Single Women by Age (1970, 1980, and 1990)

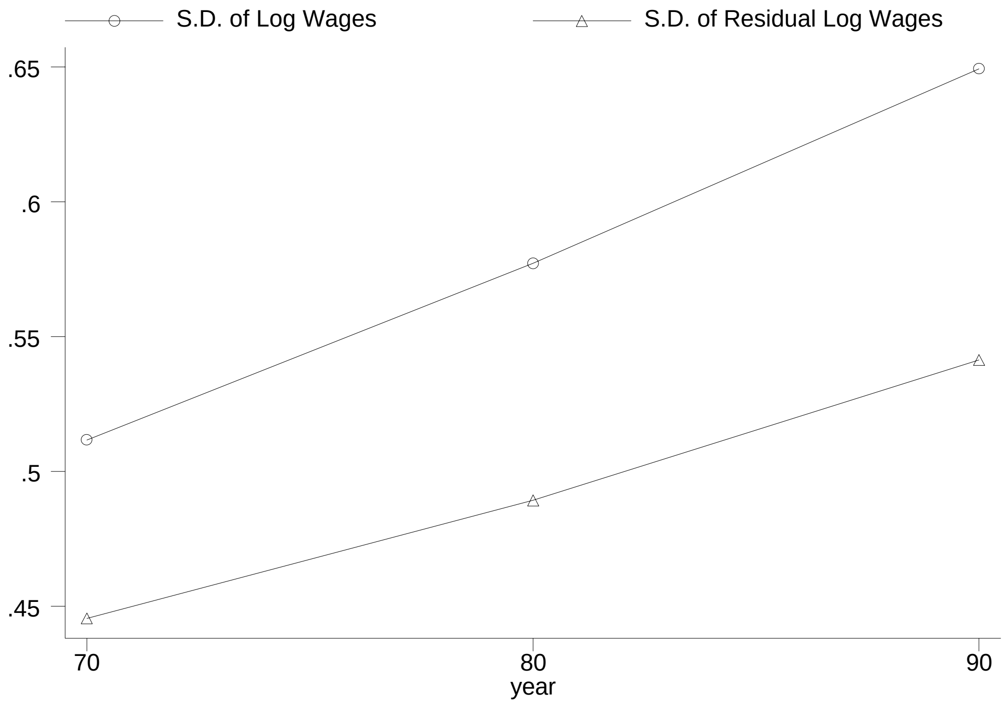

Figure 2: Wage Inequality of White Men (1970, 1980, and 1990)

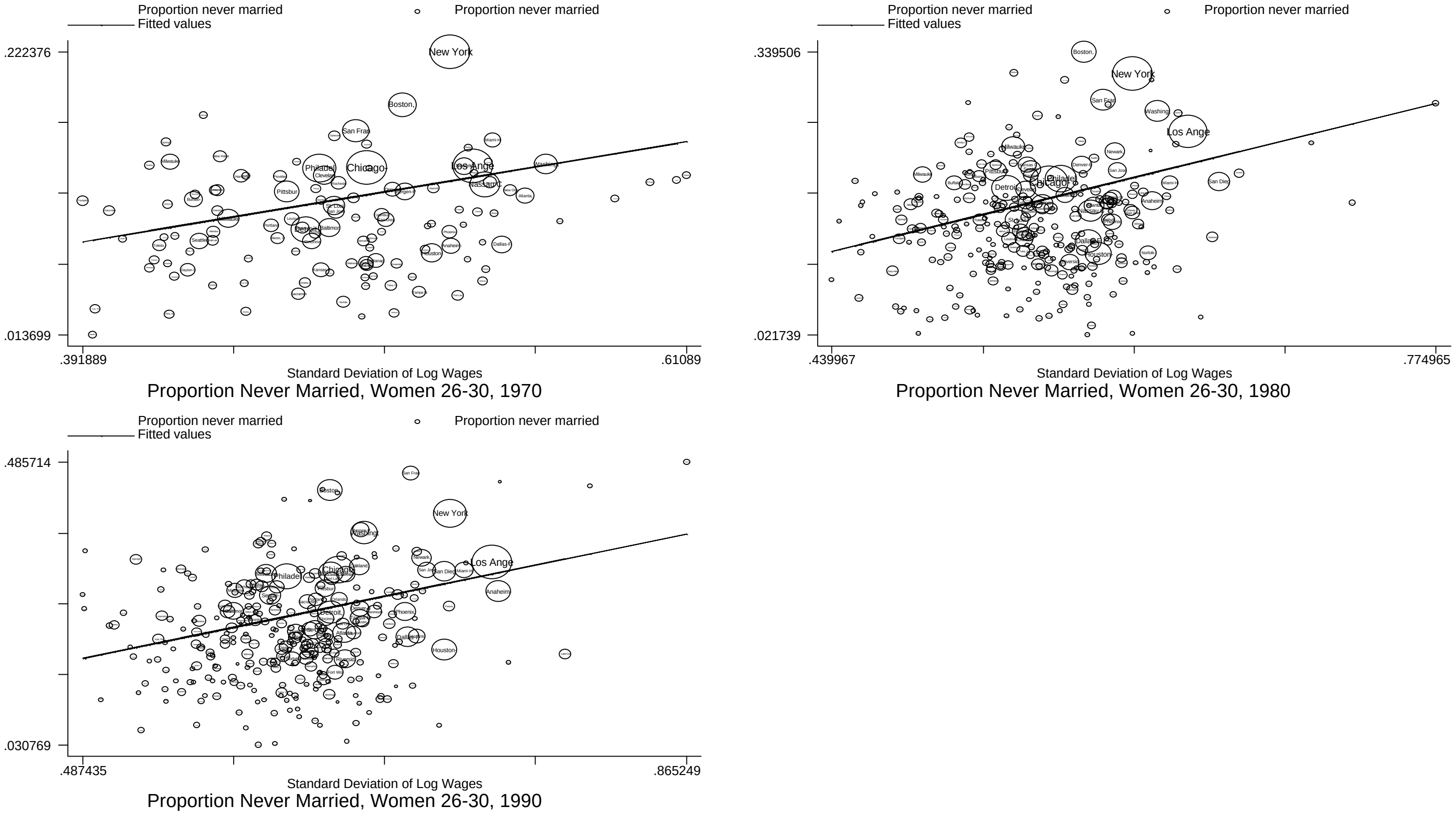

Figure 3: Marriage Rates and Inequality (1970, 1980, and 1990)

Note: The straight lines are obtained from a regression of the proportion never married on inequality by metropolitan area (weighted by population).
The regression coefficients and standard errors are:
1970: 0.217 (.095)
1980: 0.496 (.083)
1990: 0.530 (.081).

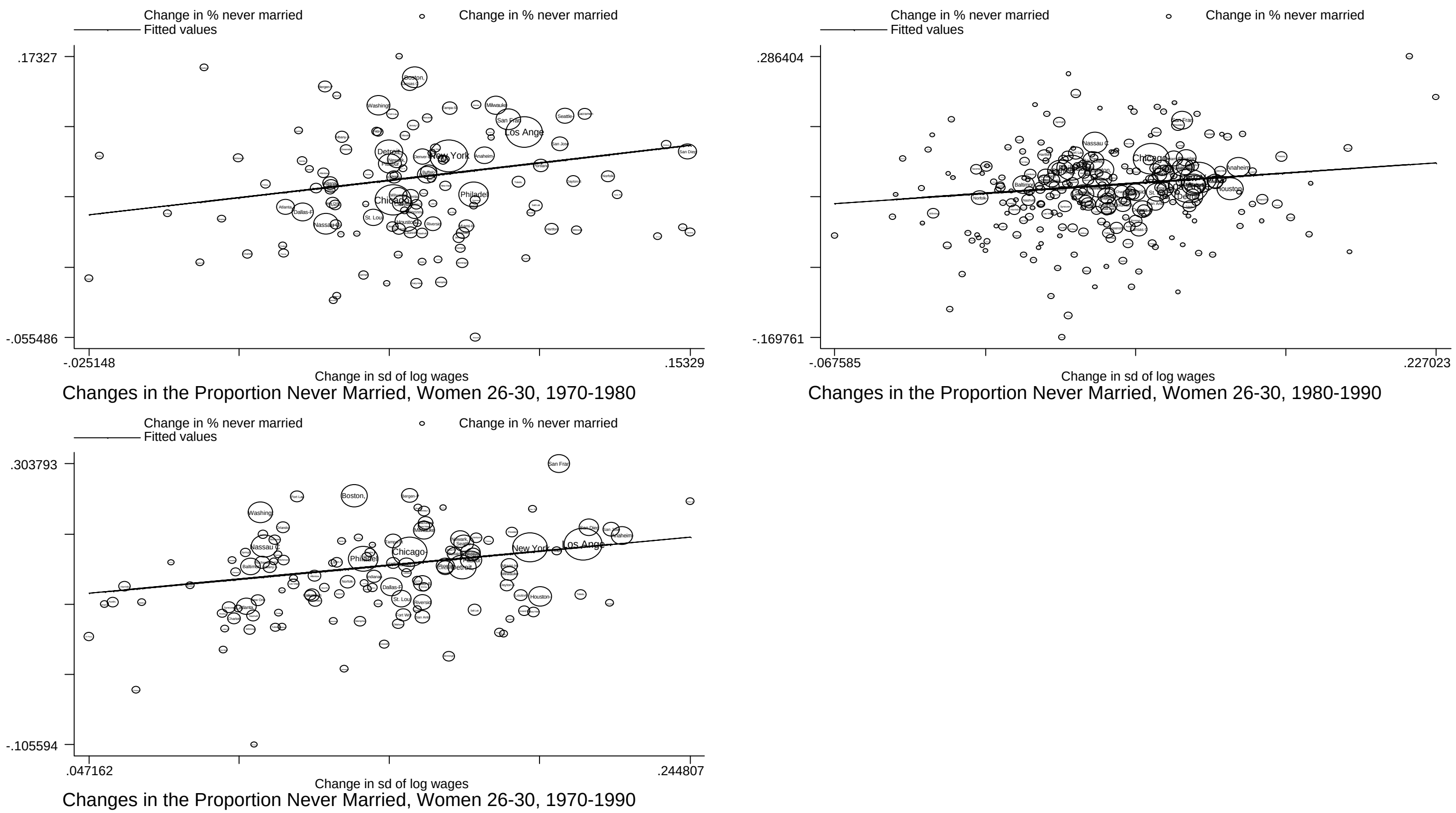

Figure 4: Changes in Marriage Rates and Inequality

Note: The straight lines are obtained from a regression of the change in the proportion never married on the change in inequality by metropolitan area (weighted by population). The regression coefficients and standard errors are:

1970-1980: 0.254 (.129)
1980-1990: 0.224 (.081)
1970-1990: 0.383 (.123).

Table 1: Sample Statistics (United States Censuses 1970, 1980, and 1990)

	1970	**1980**	**1990**
Percent White Women Age 20 Never Married	0.59	0.70	0.80
Percent White Women Age 25 Never Married	0.17	0.29	0.40
Percent White Women Age 30 Never Married	0.09	0.13	0.19
Standard Deviation of Male Log Wages	0.49	0.57	0.64
Standard Deviation of Male Residual Inequality	0.42	0.47	0.53
Mean Male Hourly Wage	15.49	15.05	14.67
Mean Female Hourly Wage	9.81	9.42	10.53
Sex Ratio (Males/Females)	0.94	0.97	0.99
Index of Relative Demand for Female Labor	0.316	0.322	0.338
Log Population	13.62	13.33	13.33

Sample includes white women between the ages of 16 and 35 who lived in a metropolitan area in the Census years 1970, 1980, and 1990. Nominal variables were converted to 1990 real values using the CPI.

Table 2: Probit for Being Single, White Women Ages 21-30 (1970, 1980, 1990)

	Pooled	**MA fixed effects**	**MA Fixed Effects + MA Linear Trend**
Standard Deviation of Male Log Wages in Metro Area	0.198 (.058)	0.273 (.068)	0.310 (.077)
Mean Male Wage in Metro Area	-0.005 (.002)	-0.008 (.003)	0.0002 (.003)
Mean Female Wage in MA	0.016 (.005)	0.014 (.003)	0.004 (.004)
Relative Demand Index for Female Labor in MA	1.227 (.122)	2.703 (.700)	4.720 (1.263)
Sex ratio in Metro Area	-0.268 (.035)	-0.139 (.055)	-0.156 (.089)
Log Population in Metro Area	0.015 (.003)	0.008 (.010)	-0.018 (.016)
HS Graduate Dummy	0.039 (.004)	0.039 (.004)	0.041 (.004)
Some College Dummy	0.184 (.006)	.183 (.006)	0.186 (.006)
College Graduate Dummy	0.305 (.006)	0.303 (.006)	0.306 (.006)
1980 Year dummy	0.117 (.007)	0.097 (.009)	0.116 (.010)
1990 Year dummy	0.167 (.011)	0.119 (.023)	0.148 (.031)
Age dummies	Yes	Yes	Yes
Geographic Fixed Effects	Region	Metro Area	Metro Area
MA Specific Linear Time Trends	No	No	Yes
No. of Metro Areas	321	321	122
No. of observations	320,788	320,788	270,242

The coefficients represent the marginal effects of the explanatory variables (evaluated at the means) on the probability of being single, estimated from a probit model. The sample includes all white women aged between 21 and 30, who lived in a metropolitan area, from the 1970, 1980 and 1990 Census years. Each specification includes an intercept. Standard errors (in parentheses) are adjusted for clustering by metropolitan area and year.

Table 3: Predicted Effects of Inequality on Marriage Trends

	Pooled	MA fixed effects	MA Fixed Effects + MA Linear Trend
	Coefficient Estimates from Table 2		
Standard Deviation of Male Log Wages in Metro Area	0.198 (.058)	0.273 (.068)	0.310 (.077)
	Predicted Effects		
% Marriage Trend Explained 1970-1990	18.4	25.4	28.8
% Marriage Trend Explained 1970-1980	15.2	21.0	23.8
% Marriage Trend Explained 1980-1990	23.1	31.9	36.2

Predicted effects are calculated by multiplying the marginal effect coefficient in Table 2 by the change in male inequality (standard deviation of weekly wages) and dividing by the average change in the marriage rate of white women ages 21-30.

Table 4: Specification and Robustness Checks

Measure of Inequality in MA	Sample Selection of White Females	Probit for Being Single: Pooled	Probit for Being Single: MA Fixed Effects
Standard Deviation of Log Male Wages	Ages 21-30	0.198 (.058)	0.273 (.068)
S.D. of Residual Male Wages (controls for age and education)	Ages 21-30	0.181 (.076)	0.220 (.080)
S.D. of Residual Male Wages (age, education, and marriage)	Ages 21-30	0.175 (.076)	0.234 (.082)
S.D. of Log Wages for Married Men	Ages 21-30	0.130 (.053)	0.208 (.071)
Gini Coefficient for Male Wages	Ages 21-30	0.306 (.069)	0.402 (.078)
90-10 Male Log Wage Difference	Ages 21-30	0.044 (.019)	0.074 (.025)
90-50 Male Log Wage Difference	Ages 21-30	0.099 (.027)	0.045 (.042)
50-10 Male Log Wage Difference	Ages 21-30	0.0005 (.027)	0.093 (.025)
S.D. of Log Male Wages	Ages 16-20	-0.035 (.039)	0.006 (.044)
S.D. of Log Male Wages	Ages 31-35	0.076 (.037)	0.128 (.043)
S.D. of Log Male Wages	Ages 21-30, education > 12	0.165 (.063)	0.253 (.090)
S.D. of Log Male Wages	Ages 21-30, education ≤ 12	0.214 (.063)	0.250 (.069)

The coefficients represent the marginal effect of the relevant inequality measure on the probability of being single, estimated from *separate* probit models. All estimating equations include an intercept and year dummies for 1980 and 1990, dummy variables corresponding to the respondent's age and completed years of education, and MA-level variables for the average male wage, average female wage, an index of relative demand for female labor, the sex ratio, and the log of total MA population. The pooled specification also includes region fixed-effects. Standard errors (in parentheses) adjust for clustering by metropolitan area and year.

Table 5: Inequality and Re-Marriage

Dependent Variable: =1 if single and divorced (or separated), =0 if divorced and remarried

	Pooled		**MA-Fixed Effects**	
	Probit	**Heckman Selection Model**	**Probit**	**Heckman Selection Model**
	Age group: 26-35			
S.D. of Male Log Wages	0.229 (.267)	0.044 (.212)	0.880 (.495)	0.756 (.564)
Marginal Effect of Probit coefficient	0.091	0.011	0.351	0.300
# Observations	40,195	179,568	40,195	179,568
	Age group: 36-45			
S.D. of Male Log Wages	0.731 (.285)	0.731 (.290)	0.729 (0.430)	0.724 (0.424)
Marginal Effect of Probit Coefficient	0.290	0.291	0.289	0.287
# Observations	35,560	140,921	35,560	140,921

Sample consists of white women from the 1970 and 1980 census years. The probit models include only women who have ever been divorced, while the Heckman specification includes all women. The first stage selection equation in the Heckman model is a probit on whether the woman ever experienced a divorce. All estimating equations include year dummies, dummy variables corresponding to the respondent's age and completed years of education, and MA-level variables for the average male wage, average female wage, an index of relative demand for male labor, the sex ratio, and the log of total MA population. Standard errors (in parentheses) adjust for clustering by metropolitan area and year.

Table 6: Effect of Female Inequality on Female Marital Status

	Probit for the Woman Being Single			
	Pooled Model		MA-Fixed Effects	
Standard Deviation of Female Log Wages in MA	-0.154 (0.080)	-0.232 (0.075)	-0.129 (0.080)	-0.182 (0.076)
Standard Deviation of Male Log Wages in MA		0.229 (0.076)		0.257 (0.077)
Geographic Fixed Effects	Regions	Regions	Metro Areas	Metro Areas

The coefficients represent the marginal effect of the relevant inequality measure on the probability of being single, estimated from *separate* probit models. All estimating equations include an intercept and year dummies for 1980 and 1990, dummy variables corresponding to the respondent's age and completed years of education, and MA-level variables for the average male wage, average female wage, an index of relative demand for female labor, the sex ratio, and the log of total MA population. Standard errors (in parentheses) adjust for clustering by metropolitan area and year.

Table 7: Does Higher Inequality Make Women Work More?

	Probit for Working Full-Time			Censored Regression (Tobit) for Weeks Worked		
Standard. Deviation of Female Residual Wages in MA	-0.330 (.086)		-0.300 (.081)	-12.133 (3.607)		-11.570 (3.442)
Standard Deviation of Male Residual Wages in MA		-0.212 (.089)	-0.158 (.083)		-5.070 (3.583)	-2.966 (3.445)
Married Dummy	-0.106 (.006)	-0.106 (.006)	-0.106 (.006)	-4.510 (0.253)	-4.510 (0.253)	-4.510 (0.253)
Number of Children Born	-0.153 (.003)	-0.153 (.003)	-0.153 (.003)	-6.156 (0.149)	-6.157 (0.149)	-6.156 (0.149)
MA Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes

Sample includes white women ages 21-30. The coefficients for the probit models represent the marginal effect of the relevant inequality measure on the probability of being single, estimated from *separate* models. The coefficients in the tobit model represent the marginal effect of the relevant inequality measure on the latent index. All estimating equations include an intercept and year dummies for 1980 and 1990, dummy variables corresponding to the respondent's age and completed years of education, and MA-level variables for the average male wage, average female wage, an index of relative demand for female labor, the sex ratio, and the log of total MA population. Standard errors (in parentheses) adjust for clustering by metropolitan area and year.

Table 8: Controlling for Women Moving Across States

	Probit for the Woman Being Single			
Standard Deviation of Male Log Wages in Current State	0.726 (.171)	0.680 (.160)	0.732 (.164)	
Standard Deviation of Male Log Wages in State of Birth				0.368 (.134)
Dummy for Current State not equal to State of Birth		-0.038 (.005)		
Current State of Residence Fixed Effects	Yes	Yes	Yes	Yes
State of Birth Fixed Effects	No	No	Yes	Yes

The coefficients represent the marginal effect of the relevant inequality measure on the probability of being single, estimated from *separate* probit models. All estimating equations include year dummies for 1980 and 1990, dummy variables corresponding to the respondent's age and completed years of education, and MA-level variables for the average male wage, average female wage, an index of relative demand for female labor, the sex ratio, and the log of total MA population. Standard errors (in parentheses) adjust for clustering by metropolitan area and year.

Table 9: The Marital Status of Men

	Probit for the Man Being Single				
Standard Deviation of Female Residual Wages in MA	-0.070 (.080)				
Standard Deviation of Male Residual Wages in MA		0.200 (.081)	0.223 (0.093)	0.181 (.094)	0.204 (.095)
Respondents' Residual Wage			-0.185 (.005)	-0.187 (.004)	0.001 (.008)
Respondent's Residual Wage Squared				0.046 (.003)	-0.003 (.0001)
Respondent's Residual Wage Rank					-0.006 (.003)
MA Fixed Effects	Yes	Yes	Yes	Yes	Yes
No. of Observations	316,252	316,252	210,851	210,851	210,851

Sample includes white men ages 21-30. The coefficients represent the marginal effect of the relevant measures on the probability of being single, estimated from *separate* probit models. All estimating equations include year dummies for 1980 and 1990, dummy variables corresponding to the respondent's age and completed years of education, and MA-level variables for the average male wage, average female wage, an index of relative demand for male labor, the sex ratio, and the log of total MA population. Standard errors (in parentheses) adjust for clustering by metropolitan area and year.